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What is the Globalisation of Inflation? ☆☆☆

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Abstract

This paper studies the globalisation of CPI inflation by analysing core, energy and food components, testing for structural breaks in the relationships between domestic inflation and a corresponding country-specific foreign inflation series at the monthly frequency for OECD countries. The iterative methodology employed separates coefficient and variance breaks, while also taking account of outliers. We find that the overall pattern of globalisation in aggregate inflation is largely driven by convergence of the mean levels of the core component from the early 1990s, compatible with the introduction of inflation targeting in many countries of our sample. There is less evidence of increased synchronisation of shortrun movements in core than aggregate inflation, but an increased role for shortrun foreign energy inflation often contributes to the globalisation effect.

1. Introduction

Over recent years policymakers and researchers have documented and discussed the globalisation of inflation, namely the apparently strong international comovement of inflation seen over the last two decades or more. Papers which document such a link between domestic and international inflation include Ciccarelli and Mojon (2010), Neely and Rapach (2011), Mumtaz and Surico (2012), Eichmeier and Pijenburg (2013), Bataa, Osborn, Sensier and van Dijk (2013), Förster and Tillman (2014). Based on this evidence, and even in the context of the large economies of the US and Euro area, Bernanke (2007) and Trichet (2008), respectively, emphasise the need for central banks to monitor carefully international price developments and analyse their implications for the domestic economy. Nevertheless, the nature of this apparent globalisation is not well understood because analyses of international inflation almost invariably employ headline or aggregate inflation.

Many heterogeneous goods and services contribute to consumer price index (CPI) inflation, but these can be usefully divided into core, energy and food. Energy and food are volatile components, with the former subject to international demand and supply shocks and the latter to the vagaries of the weather, whose effects may not persist over time. Consequently monetary policymakers often focus on inflation measures that exclude these components, conveniently referred to as core inflation; see, for

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example, the discussions in Mishkin (2007) and Bullard (2011). Although Neely and Rapach (2011) and Mumtaz and Surico (2012) argue that monetary policy plays an important role in explaining international inflation linkages, both studies focus on aggregate CPI inflation, which may disguise the effects of monetary policy changes. Further, while energy inflation has a strong international dimension, its nature may have changed since the oil price shocks of the 1970s. There have also been large changes in food supply for developed economies over the last forty years, in many cases moving from predominantly domestic production to largely imported, pointing to the possibility of increased international comovement for food inflation.

It is evident that the characteristics of aggregate inflation in developed countries has changed over the last four decades from the high levels and high volatility seen in the 1970s to low and relatively constant inflation experienced more recently. If these changes are associated with globalisation, or the process of greater integration of the world's economies, then the response of a country's domestic inflation to inflation developments in other countries must have changed at one or more points of time to make these rates more closely aligned. Bataa, Osborn, Sensier and van Dijk (2013) examine changes in short-run inflation linkages, but they consider only aggregate CPI inflation linkages for G7 countries (excluding Japan).

The present paper not only examines the role of foreign inflation for aggregate domestic CPI inflation in a range of OECD countries, but sheds new light on the nature of change by examining the three key CPI components of core, energy and food inflation[1]. The sample period extends from 1970 to 2013, with the starting date chosen to include the high inflation experience of the 1970s, which was driven at least partly by large oil price increases. To focus on comovement we employ data at the highest available frequency, namely monthly. Our sample then covers the 13 OECD countries for which aggregate and the relevant component CPI inflation data are available at the monthly frequency from 1970.

Our study is related to previous analyses, including Ciccarelli and Mojon (2010), Neely and Rapach (2011) and Mumtaz and Surico (2012), all of which examine inflation comovements in similar samples of countries to ours. Ciccarelli and Mojon (2010) document the importance of global inflation by showing that such a measure explains most of the quarterly movement in year-on-year CPI inflation since the 1960s. Employing a dynamic factor model, Neely and Rapach (2011) reinforce the importance of world and regional factors for domestic inflation, while Mumtaz and Surico (2012) extend this framework through the use of a continuously time-varying dynamic factor model. Although all of these studies discuss temporal change, their methodologies are not designed to provide formal tests for structural change in the domestic-foreign inflation relationship, which is one purpose of our analysis. Bataa, Osborn, Sensier and van Dijk (2013) is, to our knowledge, the only previous analysis of structural breaks in international inflation relationships, but their system approach limits the number of economies to be considered to three or four and also effectively assumes that breaks are coincident across countries. We use a similar methodology, but avoid the limitations of their analysis by examining each country in relation to a country-specific foreign series. Further, rather than focusing only on aggregate inflation as in almost all previous studies, we examine the roles of aggregate, core, energy and food inflation in

[1] Főrster and Tillman (2014) also study core, energy and food inflation across a similar group of countries to ours. However they consider only the post-1996 period and do not analyse structural breaks.

this international context.

Domestic-foreign inflation links are first studied separately for the four measures of interest, namely the three components and aggregate CPI inflation. More formally, we test for structural breaks in a dynamic model for domestic monthly inflation in relation to the corresponding foreign series, with the latter constructed as the bilateral trade-weighted average of the relevant inflation series in the other countries of our sample. The structural breaks uncovered point to the globalisation of aggregate CPI inflation, which (for most countries in our sample) is marked by an increased contemporaneous response of domestic to foreign inflation from the 1980s. Further analysis points to a key role played by energy inflation in this increased short-run comovement and by core inflation for the apparent convergence of the level and volatility of aggregate inflation across these economies.

The results also indicate that all economies in our sample that introduced inflation targeting in the early 1990s, experienced a decline in core or aggregate domestic inflation during that period; most of them also experienced a decline in core and aggregate inflation persistence. However, we find no clear evidence of a positive and increasing shortrun comovement in core inflation among the economies examined. This indicates that the observed convergence in aggregate and core inflation may be the product of many economies sharing a similar inflation target concurrently, rather than due to a global transmission factor.

The rest of the paper is organized as follows. Section 2 describes our data, while Section 3 discusses the methodology we employ. Substantive results are reported and discussed in the following two sections, with the nature of change documented for individual (aggregate and component) series in Section 4 and breaks in aggregate inflation decomposed in terms of foreign core, energy and food inflation series in Section 5. Finally, Section 6 concludes.

2. Data

As just discussed, the inflation data we study comprise monthly CPI aggregate inflation, together with the corresponding core, energy and food component series, for the OECD countries for which such data are available to the beginning of the 1970s. To be specific, our sample period extends from January 1970 to September 2013, thereby extending from prior to the oil price shocks of the 1970s to the post-GFC (global financial crisis) era. All data are sourced from the OECD Main Economic Indicators database. The sample includes six member countries of the Euro area (Austria, Finland, France, Germany, Italy, Netherlands), four further European countries (Denmark, Sweden, Switzerland, UK) and three other major economies (Canada, Japan, US). The sample also includes four inflation targeting economies (Canada, UK, Sweden and Finland)[2]. Although selected purely on data availability, the inclusion of a number of Euro area and inflation targeting countries may shed light on the impact of these two economic innovations on the nature of inflation in these countries.

It should be noted that the inflation series employed in our analysis are not entirely comparable across countries and sometimes over time[3]. The most important caveat in this respect is the differing treatments of owner-occupied housing in aggregate and core inflation. Some countries, such as the US

[2]Finland introduced inflation targeting in early 1993, following an earlier announcement, but abandoned it on joining the Eurozone in January 1999

[3]We thank the referee for drawing our attention to this important point.

and Japan include this as imputed rent (with substantial weight for both countries), Canada employs a user-cost approach, while other countries, including France and the UK, exclude owner-occupied housing from CPI. Further, although restaurant meals are generally excluded from the food index, this is included in the UK series before 1985. Despite these comparability issues, we are confident that the compositions of the various series are sufficiently consistent for us to be able to draw substantive conclusions relating to the globalisation of inflation.

All monthly inflation series are calculated as 100 times the one month change in the log index. Prior to analysis, each series is seasonally adjusted using the widely applied X-12-ARIMA procedure[4]. We employ monthly inflation rates in preference to annual inflation computed on a monthly basis for statistical reasons. In particular, the smoothing inherent in the computation of annual differences leads to adjacent observations overlapping by eleven months, thereby inducing a long moving average component in the disturbances. The resulting moving averages may be poorly approximated by the autoregressive lag structure used in modelling and, further, the more highly parameterised models then required would be expected to reduce the power of the structural break tests that are essential to our analysis.

Although our formal analysis employs month-on-month inflation, volatility in such series largely obscures other patterns when the data are examined visually. Therefore, Figure 1 shows percentage annual inflation in our sample of countries, for each of aggregate, core, energy and food CPI inflation. The general decline over time in both the level and volatility of aggregate inflation is clear from panel a, while the relatively low cross-country dispersion over the last two decades is also evident. Since the weights attached to the energy and food components are relatively low compared with core inflation, the patterns seen in core inflation in panel b largely reproduce those for the aggregate series in panel a. Although the level of food inflation (panel d) is also generally lower from around 1990 than previously, there appears to be greater cross-country dispersion in this latter period than in panel a or b. Volatility in food inflation is also higher than for aggregate or core inflation. It is well known that energy inflation is volatile over time, but panel c also emphasises the communality of movements, at least in annual inflation as plotted here. Therefore, Figure 1 makes clear that the three component series have distinctive properties, underlying the importance of a formal examination of their roles in the (apparent) globalisation of aggregate inflation.

2.1. Foreign inflation

Dynamic factor model studies, including Neely and Rapach (2011), Mumtaz and Surico (2012), and Főrster and Tillman (2014), represent foreign (or world) inflation as a common factor extracted from a cross-country range of inflation series. Ciccarelli and Mojon (2010), on the other hand, focus on a global series computed as an average across the OECD countries in their sample. Ciccarelli and Mojon (2010) also consider the use of aggregated OECD CPI inflation and an extracted common factor, finding the resulting domestic-global inflation relationships to be largely unaffected by this choice.

All the global inflation measures used in the studies just mentioned are constructed from data that includes the country whose domestic inflation is being considered. This, therefore, induces some

[4] Official seasonally adjusted CPI data are available for the US and a graphical comparison of this series with the comparable unadjusted series filtered using X-12-ARIMA showed these to have very similar properties. Hence we apply X-12-ARIMA seasonal adjustment to series for all countries.

simultaneity in the domestic-global relationship. Although the effect may be small when global inflation is computed over a large number of countries, we nevertheless prefer to capture international influences on domestic inflation through the use of a country-specific foreign series which is constructed excluding the country of interest.

Therefore, we compute foreign inflation for country s, where $s = 1, ..., N$ and N is the number of countries analysed, as a weighted average of the corresponding inflation series over the other $N-1$ countries in the data set. Empirical open economy models of inflation often employ trade weights to construct relevant foreign series; see, for example, Ihlig, Kamin, Lindner and Marquez (2010) and the references therein. Following this approach, our main results are based on foreign inflation for country s computed from bilateral trade statistics as[5]

$$w_{t,s}^{(i)} = \frac{\psi_{t,s}^{(i)}}{\sum_{i=1,i\neq s}^{N} \psi_{t,s}^{(i)}} \tag{1}$$

$$\text{where} \qquad \psi_{t,s}^{(i)} = (M_{t,s}^{(i)} + X_{t,s}^{(i)}) \tag{2}$$

$$\text{and} \qquad \sum_{i=1,i\neq s}^{N} w_{t,s}^{(i)} = 1. \tag{3}$$

That is, the trade weight for country s with respect to country i for month t, $w_{t,s}^{(i)}$, is measured by the sum of imports from i ($M_{t,s}^{(i)}$) and exports to i ($X_{t,s}^{(i)}$) as a proportion of all trade that month for country s with the $N = 13$ countries we study. Although the weights are time varying and computed for each individual month, changes over time are generally small, implying the trade structure across the countries in our sample changes only gradually. After computing trade weights, country specific foreign inflation is constructed as

$$\pi_{t,s}^{F} = \sum_{i=1,i\neq s}^{N} w_{t,s}^{(i)} \pi_{t,i}. \tag{4}$$

The same computations, with weights as in equations (1) to (4), are used to construct both aggregate foreign inflation and the core, energy and fuel sub-aggregates.

Data on bilateral imports and exports used in equation (2) are also obtained from the OECD Main Economic Indicator database, with all series expressed in US dollars using (where appropriate) the irrevocable exchange rates to account for the establishment of the Euro area. In general, movements in exchange rates affect trading quantities, which in turn affect imported prices and eventually inflation. Similarly, exchange rate regimes matter for inflation and this is why their role is discussed in studies that use earlier data that cover part of the Bretton Woods period, (Ciccarelli and Mojon, 2010, Neely and Rapach, 2011 and Mumtaz and Surico, 2012). Exchange rates are, however, much more volatile than inflation and so to keep their models tractable, even these studies of global inflation do not explicitly account for the effects of movements in exchange rates. In this respect our approach is similar, however there are two notable differences. First, our data sample starts in the 1970s, thus excluding the Bretton Woods period. Second, the weights we use to compute foreign inflation are time varying, capturing trade changes over time that may be partly due to exchange rate movements.

[5]This definition corresponds to that used in Cesa-Bianci *et al.* (2012).

Table A.1 of the Appendix shows the bilateral trade weights averaged over our sample of approximately 40 years. In general, Germany is the most important trading partner for European countries, while the US plays this role for the non-European countries of Japan and Canada. However, the UK does not have a dominant trading partner, although shares with respect to the US, Germany and France are relatively large compared to others. These weights based on bilateral trade do not reflect effects of other countries, such as the large emerging economies of China and India, but the limitation of available data for those countries precludes their inclusion in our analysis.

Using the bilateral trade variable (as in equation (2)) carries the potential for the resulting international inflation variable to be endogenous in a model for inflation[6]. It can be argued that foreign and domestic inflation will have a direct effect on a country's imports and exports. For this reason we examine whether the main results of this paper are robust to changes in the definition of the weighting variable, $\psi_{t,s}^{(i)}$. In particular we choose two alternative specifications that arguably significantly reduce the potential endogeneity problem. First we set the weighting variable to be the total trade of the respective partner country, i, rather than merely the bilateral trade, $\psi_{t,s}^{(i)} = \sum_{j=1,j\neq s}^{N}(M_{t,i}^{(j)} + X_{t,i}^{(j)})$. When calculating total trade we exclude the trade with country s, the country for which we will formulate an inflation model. Lastly we set the weighting variable to the partner countries' GDP, $\psi_{t,s}^{(i)} = GDP_{t,i}$.

While it was apparent from Table A.1 that the weights calculated on the basis of bilateral trade differed substantially between countries s, the weights calculated on the basis of the alternative total trade and GDP variables largely remain unchanged between countries as they by and large reflect the size of the partner country i[7].

In order to understand the impact of these different definitions we display the resulting foreign inflation series for the US in Figure 2. Differences arise mainly in the volatility of the resulting series. The foreign inflation variable that is based on the use of the bilateral weights in equation (2) results in a more variable foreign inflation series when compared to those calculated using the total trade or GDP weighting variable[8].

3. Methodology

This section turns to the methodology we use to detect structural breaks in the domestic-foreign inflation relationships for each of the 13 countries we study. Although there is a substantial empirical literature covering many countries that examines changes in the univariate properties of inflation (see, for example, Cecchetti and Debelle, 2006, or Altissimo *et al.*, 2006), relatively little attention has been paid to the implications of variations in inflation volatility in this context. This is a serious limitation, since Pitakaris (2004) shows that inference on coefficient breaks is distorted in the presence of volatility changes. However, Pitakaris (2004) also provides a solution, showing that variance breaks can be satisfactorily taken into account through a generalised least squares (GLS) transformation.

[6]We thank an anonymous referee for highlighting this issue.

[7]Tables characterising the resulting weights are not shown here to conserve space but are available from a Web Appendix: `https://dl.dropboxusercontent.com/u/35924101/ABBO2016WebAppendix.docx`

[8]As it turns out the US foreign inflation series in Figure 2 displays the most pronounced differences. The differences for the other countries are even less obvious. Equivalent Figures for the remaining countries can be found in the Web Appendix.

Following and building upon Bataa, Osborn, Sensier and van Dijk (2013, 2014), our approach identifies breaks in the coefficients and variance by iterating between tests for these two types of breaks. Variance breaks uncovered are removed through a GLS transformation when testing for coefficient breaks. Further, since ignoring outliers in the data can lead to model misspecification biases in the estimated parameters (see, among others, Giordani, Kohn and van Dijk, 2007, or Chen and Liu, 1993), our procedure also takes account of outliers.

Subsections 3.1 to 3.3 below describe the model and methodology used for examining the domestic-foreign inflation linkages for each of aggregate CPI and its core, energy and food components. Subsection 3.4 then generalises the approach to consider the relationship between domestic aggregate inflation and foreign component series.

3.1. The domestic-foreign inflation model

Our model relates domestic inflation for country s ($s = 1, ..., N$) in month t ($\pi^D_{t,s}$) to a corresponding measure of foreign inflation ($\pi^F_{t,s}$); the inflation series under analysis can be either aggregate CPI inflation or a component (core, food or energy). In a dynamic context, the relationship can be parsimoniously represented as

$$\pi^D_{t,s} = \alpha_0 + \sum_{i=1}^{p} \alpha_{i,s}\pi^D_{t-i,s} + \beta_{0,s}\pi^F_{t,s} + \sum_{i=1}^{r} \beta_{i,s}\pi^F_{t-i,s} + \varepsilon_{t,s}, \quad t = 1, ..., T \tag{5}$$

where $\pi^F_{t,s}$ is foreign inflation for country s at time t, $\beta_{0,s}$ and $(\beta_{1,s}, ..., \beta_{r,s})$ capture the contemporaneous comovement and spillovers (respectively) between domestic and foreign inflation, $(\alpha_{1,s}, ..., \alpha_{p,s})$ are own (domestic) inflation dynamics, α_0 is an intercept, $\varepsilon_{t,s}$ is a temporally uncorrelated disturbance process and T observations are available for estimation, after allowing for lags. Structural break tests are employed to examine whether the coefficients of equation (5), including the disturbance variance $E[\varepsilon^2_{t,s}] = \sigma^2_s$ are constant over time.

The motivation for our structural break testing is that the globalisation of inflation implies that the process for $\pi^D_{t,s}$ has experienced one or more changes over time, such that it has become more strongly influenced by inflation in other countries. Thus, we examine whether the coefficients associated with $\pi^F_{t,s}$ have increased, particularly the contemporaneous coefficient $\beta_{0,s}$. From a long-run perspective, the mean levels of domestic and foreign inflation will become more closely aligned. Persistence and volatility properties may also be expected to become more internationally aligned as inflation globalisation proceeds. Therefore, we examine both the coefficients and disturbance variance of equation (5).

With structural breaks in the coefficients of equation (5) at m dates $T^C_1, ..., T^C_m$, then there are $m + 1$ coefficient regimes and we can write[9]

$$\pi^D_{t,s} = \alpha_{0j} + \sum_{i=1}^{p} \alpha_{ij,s}\pi^D_{t-i,s} + \beta_{0j,s}\pi^F_{t,s} + \sum_{i=1}^{r} \beta_{ij,s}\pi^F_{t-i,s} + \varepsilon_{t,s}, \quad t = T^C_{j-1}+1, ..., T^C_j,\ j = 1, ...m+1, \tag{6}$$

with the convention that $T^C_0 = 0$ and $T^C_{m+1} = T$. Within each coefficient regime j for country s, the

[9]The number of coefficient regimes can vary over countries, but for notational simplicity this is denoted simply as m rather than m_s. Similarly, the break dates are country-specific.

coefficients $\alpha_{j,s} = (\alpha_{0j,s},\, \alpha_{1j,s},\, ..., \alpha_{pj,s})'$ and $\beta_{j,s} = (\beta_{0j,s},\, \beta_{1j,s},\, ..., \beta_{rj,s})'$ are time-invariant and all autoregressive (AR) roots are assumed to lie strictly outside the unit circle. With n volatility breaks, the innovation variance $E[\varepsilon_{t,s}^2] = \sigma_{k,s}^2$ is homoscedastic within each volatility regime, but varies over the $n+1$ volatility regimes $k = 1, ..., n+1$. The numbers of breaks (m and/or n) may be zero and dates of all breaks are unknown. No restrictions are imposed across coefficient and volatility breaks, so that these may differ in both numbers of breaks and their timing.

The maximum lag order considered for both own and foreign inflation is $p = r = 12$, but the included lags are specified using a general to specific approach, as explained below. In line with the usual definition employed in a univariate context, $\rho_{j,s}^d = \sum_{i=1}^{p} \alpha_{ij,s}$ is referred to as inflation persistence (in our case, for country s in regime j), although it is measured in equation (6) conditionally on foreign inflation.

Although excluded from equations (5) and (6) for simplicity of representation, our procedure allows for the presence of additive outliers in $\pi_{t,s}^D$, which could be due to (say) changes in indirect taxes. As explained in the next subsection, consideration of outliers allows for both coefficient and volatility breaks in equation (6).

3.2. Iterative testing methodology

The iterative methodology proposed by Bataa *et al.* (2014) employs structural break tests in conjunction with the outlier detection and removal procedure of Stock and Watson (2003) to examine structural breaks in each of the seasonal, mean, dynamic and volatility components of univariate inflation. While based on this approach, our procedure differs in three important respects. Firstly, we treat all elements of the regime j coefficient vector $\delta_{j,s} = (\alpha_{j,s}', \beta_{j,s}')'$ of equation (6) together, rather than separating mean and dynamic breaks. This is primarily because the simulation analysis in Bataa *et al.* (2014, Table 1) indicates that their procedure yields a mean break test that is oversized. This feature may be due to the initial tests for mean breaks applying HAC (heteroscedasticity and autocorrelation consistent) inference with unmodelled dynamics, as Bai and Perron (2006) also show that such inference can lead to badly oversized tests. However, Bai and Perron (2006) show that inference is improved when the dynamics are explicitly modelled, as in our approach. Secondly, our procedure simplifies the iterations in respect of volatility breaks, since the results of Bataa *et al.* (2013, Table 1) implies these are detected well without iteration. Finally, we take account of variance breaks when identifying outliers, whereas the methodology of Bataa *et al.* (2014) does not[10].

Unlike Bataa *et al.* (2014), seasonality is not of interest in the present study and (as discussed in Section 2) all series are seasonally adjusted prior to the application of the iterative procedure. This is justified by the robustness analysis in Bataa *et al.* (2014), which indicates that the detection of structural breaks in the mean and dynamics is not substantively affected by the method of accounting for seasonality. Our iterative testing methodology involves five steps which are outlined below and explained in more detail in the Appendix.

Step 1: Outlier detection. Outliers in $\pi_{t,s}^D$ are detected and replaced using the procedure of

[10] To be more explicit, we take account of breaks in each domestic inflation series. A small Monte Carlo analysis confirmed that the presence of aberrant observations in the explanatory variable did not affect the size of the structural break tests for equation (6).

Stock and Watson (2003), rescaling observations to take account of volatility breaks and treating each coefficient regime separately (both regimes as detected in the previous iteration)[11].

Step 2: Model selection. After correcting for outliers, the dynamics of the model in equation (6) are specified using a general to specific procedure in combination with the Schwartz Information Criterion (SIC).

Step 3: Preliminary coefficient break test. The Bai and Perron (1998) HC multiple structural breaks test procedure is applied to the coefficients of the model in equation (6).

Step 4: Variance break test. Variance breaks are detected in the residuals of equation (6) allowing for coefficient breaks identified in step 3.

Step 5: Coefficient break test. Coefficient breaks are reconsidered, taking account of variance breaks detected in step 4 by applying the feasible GLS transformation.

A single iteration comprises steps 1 to 5, inclusive, and the output of each iteration consists of coefficient and volatility break dates, together with an outlier-corrected series. The maximum number of iterations is set to 10 and convergence may be achieved in either of two ways. Firstly, identical dates may be detected in two consecutive iterations; alternately, the iterations may cycle between (say) two or three sets of dates. In the latter case, we focus on coefficient breaks and choose the set which achieves the smallest SIC among those in the cycle. The version of SIC is that proposed by Yao (1988) for structural break inference, which is applied to the GLS transformed data and calculated for m coefficient breaks as

$$SIC(m) = \ln[T^{-1}S_T(\hat{T}_1^C, ..., \hat{T}_m^C)] + q^* \ln(T)/T \tag{7}$$

where $S_T(\hat{T}_1^c, ..., \hat{T}_m^c)$ is the sum of squared standardised residuals for $\pi_{t,s}^D$ computed over the $m+1$ coefficient regimes in equation (6) and $q^* = (m+1)q + m$ where q is the total number of coefficients (including the intercept) estimated in the model. Note that, through q^*, the penalty term effectively treats each coefficient break date as an estimated parameter. The T sample observations for $\pi_{t,s}^D$ used in computing SIC are identical over all models in the comparison.

3.3. Break inference

The heart of the iterations described in subsection 3.2 is the multiple structural break testing procedure of Bai and Perron (1998). To be more specific, and assuming m coefficient breaks, the optimal break dates and corresponding coefficient estimates are computed using the dynamic programming algorithm of Bai and Perron (1998, 2003a), which searches for the minimum total residual sum of squares over all $m+1$ regimes. However, the true m is unknown and a specified maximum number of break dates to be considered (say m^*) leads to the identification of m^* sets of possible estimated break dates, namely for $m = 1, 2, ..., m^*$ breaks.

Prior to selecting a specific number of breaks from these possible break date sets, and as recommended by Bai and Perron (2003a), a preliminary test of the overall null hypothesis $H_0 : m = 0$ is undertaken against the composite alternative $H_A : m = 1, ..., m^*$. The statistic $WDmax$ is used for

[11] In the initial iteration, outliers are judged over the whole sample with no rescaling.

this purpose[12], and failure to reject H_0 implies no breaks are found to occur. Also as recommended by Bai and Perron (1998, 2003a), when H_0 is rejected, their sequential $SupF\,(l+1|l)$ test is employed to estimate the appropriate number of breaks, $\hat{m}$. That is, the null hypotheses of $l = 1, 2, 3, ...$breaks (subject to a maximum of m^* breaks) are examined sequentially against the alternative of $l+1$ breaks, with the first non-rejection yielding $\hat{m} = l$. (The null of $l = 0$ is not considered against $l = 1$, since the overall null has been rejected.) All tests are computed at a nominal 5 percent level of significance, with a maximum of $m^* = 5$ breaks considered. Testing employs the asymptotic distributions obtained by Hall and Sakkas (2013), which these authors show to be more accurate than the critical values provided by Bai and Perron (2003b) and have the additional advantage of allowing computation of approximate asymptotic p-values. The so-called trimming parameter, which defines the minimum distance between two consecutive breaks as a function of the total sample size T, is set to 0.15.

The volatility break tests are implemented as follows. Using the $\hat{m}$ initial coefficient break date estimates of step 3 (subsection 3.2) and the associated regime-specific estimated coefficients, residuals $e_{t,s}$ for country s are computed for equation (6) over $t = 1, ..., T$. These are employed in the test regression

$$e_{t,s}^2 = \gamma_{j,s} + u_{t,s} \tag{8}$$

where $\gamma_{j,s}$ is constant within the volatility regime j, but is allowed to change over time due to the presence of volatility breaks. The same procedure as just described for the coefficients is applied to obtain the estimated number of volatility breaks ($\hat{n}$) and the associated volatility break dates $(\hat{T}_1^V, ..., \hat{T}_{\hat{n}}^V)$. When breaks are uncovered, the standard error in each volatility regime $j = 1, ..., \hat{n}+1$ is calculated as

$$\hat{\sigma}_j = \sqrt{(\hat{T}_j^V - \hat{T}_{j-1}^V)^{-1} \sum_{t=\hat{T}_{j-1}^V+1}^{\hat{T}_j^V} e_{t,s}^2},$$

which is used to apply the GLS transformation to all variables of equation (6) for the coefficient break tests of step 5.

3.4. Decomposing foreign inflation

To shed further light on the nature of changes in international inflation linkages, or (possible) globalisation, we also study a generalised version of the model in equation (6), namely

$$\pi_t^D = \alpha_{0j} + \sum_{i=1}^{p} \alpha_{i,j}\pi_{t-i}^D + \sum_{k=1}^{3} \left\{ \beta_{0k,j}\pi_{t,k}^F + \sum_{i=1}^{r} \beta_{ik,j}\pi_{t-i,k}^F \right\} + \varepsilon_t, \;\; t = T_{j-1}^C+1, ..., T_j^C, \;\; j = 1, ..., m+1 \tag{9}$$

where π_t^D is domestic aggregate inflation and $\pi_{t,k}^F$ ($k = 1, 2, 3$) are the component foreign series relating to core, energy and food inflation in month t, and volatility breaks are also permitted. Although the domestic and foreign inflation series, together with the respective coefficients, are specific to country s, this subscript is dropped from equation (9) for notational simplicity.

[12] The $WDmax$ statistic is used in preference to $UDmax$ because it embodies a set of weights that ensure the marginal p-values are equal for the null of $m = 0$ against each specific number of breaks $m = 1, ..., m^*$ (Bai and Perron, 1998).

The generalised model in equation (9) has more parameters than that in equation (6) for aggregate inflation. Due to the costs of searching in this more highly parameterised model, the dynamics are not respecified within each iteration. Specifically, the dynamic specification is selected prior to the commencement of the iterations using the same model selection procedure as explained in the Appendix for equation (6). This set of lags is then used throughout the iterative testing procedure (essentially omitting Step 2 described in Subsection 3.2).

Due to model (9) being highly parameterised it is possible that the break identification strategy may lack power. For this reason we employ an alternative break identification strategy for model (9) allowing a robustness check of our findings. In this alternative strategy we employ the outliers, lag specification and break dates identified using the iterative procedure applied to the aggregate inflation model (6). The specification for the aggregate foreign inflation lags that is found in that model is then directly applied to the three foreign component series without any further specification testing.

The use of equation (9) in place of equation (6) effectively treats aggregate foreign inflation as a weighted sum of the relevant component series with constant weights. However, the weights change over time, both those we use to construct each foreign inflation series from data for other countries (see subsection 2.1) and those used by the national statistical agency in constructing each country's domestic aggregate inflation. Hence it is only an approximation to consider aggregate foreign inflation as a fixed weighted sum of the corresponding component series. This is, indeed, an additional reason why it is appropriate to re-evaluate the existence and dates of breaks in international inflation linkages in the context of model in equation (9).

Reflecting our primary focus on the nature of changes in the international comovement of inflation, we apply a sequence of conventional F-tests to the model in equation (9)[13]. The contemporaneous coefficients may be considered particularly relevant to globalisation and hence we test the null hypothesis of no change across regimes

$$H_0 : \beta_{0k,1} = ... = \beta_{0k,m+1} \tag{10}$$

separately for each component $k = 1, 2, 3$ and also jointly across all three components

$$H_0 : \beta_{0k,1} = ... = \beta_{0k,m+1}, \quad \text{all } k = 1, 2, 3. \tag{11}$$

The tests of equations (10) and (11) are applied to both sets of estimated break dates for equation (9), namely those based on the aggregate CPI model of (6) and those estimated directly from equation (9). In addition, the corresponding tests are applied to the lagged coefficients for foreign inflation, namely

$$H_0 : \beta_{ik,1} = ... = \beta_{ik,m+1}, \text{ all } i = 1, ..., r \tag{12}$$

and

$$H_0 : \beta_{ik,1} = ... = \beta_{ik,m+1}, \quad \text{all } i = 1, ..., r \text{ and } k = 1, 2, 3. \tag{13}$$

Although it is tempting to compare the coefficients of foreign inflation components estimated in

[13] Bai and Perron (1998) show that (with coefficient breaks of fixed magnitude) the estimated break fractions asymptotically converge to the true values at a rate of T, whereas the estimated coefficients in a model such as equation (9) converge at the rate of $T^{1/2}$. This implies that conventional hypothesis tests applied to the coefficients are asymptotically valid when conditioned on the estimated break dates.

equation (9) with those for the separate models estimated using equation (6), care must be taken when doing so due to the weighting involved in equation (9). To see this, consider the following special case of equation (6) for each inflation component k

$$\pi_{tk}^D = \alpha_0 + \alpha_1 \pi_{t-1,k}^D + \beta_0 \pi_{tk}^F + \varepsilon_{tk}, \quad k = 1, 2, 3 \tag{14}$$

where no structural breaks apply and the country subscript s is omitted for simplicity. Note, in particular, that equation (14) assumes common coefficients across the three components. Also assume the components (core, energy and food) have constant weights in forming the aggregate series, both over time and across domestic and foreign inflation, with $\pi_t^D = \sum_{k=1}^3 \omega_k \pi_{tk}^D$ and $\pi_t^F = \sum_{k=1}^3 \omega_k \pi_{tk}^F$. Aggregating equation (14) across components gives

$$\pi_t^D = \alpha_0 + \alpha_1 \pi_{t-1}^D + \sum_{k=1}^{3} (\omega_k \beta_0) \pi_{tk}^F + \varepsilon_t. \tag{15}$$

Consequently, the contemporaneous coefficient of the foreign component series k in equation (15) is not β_0, but rather $\beta_{0k} = \omega_k \beta_0$.

Consideration of this simple special case implies that coefficients estimated in the component models in equation (6) should be scaled by the weights ω_k when compared with those estimated in equation (15). In the case of the US, for example, over the period 1987 to 2012, goods and services that contribute to core inflation have an average weight of 0.77 in aggregate inflation, with those for energy and food inflation being 0.09 and 0.14, respectively. Therefore, the estimated coefficients of foreign energy and food inflation can be anticipated to be substantially smaller in the context of equation (9) than in the separate models of equation (6), even when the same coefficients apply in the latter across the three components. Due to the larger role it plays in the aggregate, the reduction will be less marked for core inflation.

4. Globalisation of Inflation Characteristics

Figures 4 to 8 visually present the results of our analysis of the globalisation of inflation through the model in equation (6), with further details given in Table 1 and Appendix Tables A.2 to A.5. These results are set in context by subsection 4.1, which discusses the model specification and the evidence of breaks in the coefficients and residual variances of equation (6) including robustness to the choice of weights in the construction of foreign inflation. Subsections 4.2 and 4.3 then focus on the evidence of globalisation seen in the coefficients of contemporaneous foreign inflation and the mean inflation levels, respectively. Changes in persistence and volatility characteristics of inflation are considered in subsection 4.4, which is followed by a discussion of the implications of our findings.

4.1. Preliminary results

In a first step we discuss the nature of the dynamic specification of equation (6) as selected within our iterative structural break testing procedure (see Section 3) with foreign inflation constructed using bilateral trade weights (Section 2.1). In Table 1 we display the lags of the respective domestic and foreign inflation series that have been selected (maximum lag $p = r = 12$). For example, in the model for Canadian aggregate inflation we find lags 5, 7 and 9 of the domestic aggregate inflation to be statistically significant as well as the contemporaneous and 9 months lagged foreign aggregate inflation series. The columns labelled (Aggregate, Core, Energy and Food Inflation) display the selected lags for all countries considered for the models of the respective inflation series[14].

Reflecting our focus on inflation linkages, contemporaneous foreign inflation is included in all models used for structural break analysis, even when the dynamic specification procedure would omit this; such cases are indicated by 0^* in Table 1. The domestic aggregate-foreign components model of Table 1 refers to the component model (equation (9)), which is discussed in the next section.

The results in Table 1 illustrate that the models for energy inflation have very simple dynamics, sometimes (Finland, Germany, Sweden and Switzerland) domestic energy inflation reacting to contemporaneous foreign inflation only. The selected lag structures for the models for aggregate, core and food inflation tend to be more complex, in particular for the lags of the domestic inflation series. Core inflation models tend to have the most complex domestic lag structures, but in turn frequently exclude contemporaneous foreign inflation (eight of the 13 countries). Both food and energy inflation react very quickly to foreign inflationary pressures, reflected in seven countries requiring only contemporaneous foreign inflation in the food and energy inflation models.

These results provide little evidence that globalisation applies to shortrun monthly movements in core inflation. This finding is in line with that reached by Förster and Tillman (2014) for the post-1996 period. The fact that the models for aggregate inflation always select the contemporaneous foreign aggregate inflation lag can therefore be ascribed to the influence of the food and energy components of aggregate inflation. Stated differently, the results in Table 1 provide a first indication that international linkages for core inflation are distinctly different to those in aggregate CPI inflation. This finding extends previous analyses (such as Nason, 2006) that find different properties for these series in a domestic US context.

Details of the breaks uncovered in the coefficients and disturbance variance of equation (6) for each inflation series can be found in Appendix Table A.2. The results labelled aggregate-foreign components in that table relate to the discussion of Section 5. Further, Appendix Tables A.3 and A.4 show detailed results for both the overall $WDmax$ statistic and sequential F-tests for equation (6) applied to the aggregate inflation series in the final iteration of the procedure of Section 3. These latter tables include not only the test statistic values, but also the corresponding p-values obtained using the method of Hall and Sakkas (2012)[15]. The corresponding results for core, energy and food inflation models are not shown in order to conserve space, but are available from the authors on request. As illustrated for aggregate inflation in Tables A.3 and A.4, the $WDmax$ test p-values provide strong evidence for breaks in both the coefficients and disturbance variances of inflation. However, these tables also indicate that the number of breaks is not always clear-cut, with some sequential tests resulting in marginal p-values in relation to a 5% significance level.

To summarise the results, based on a 5% significance level throughout, coefficient and volatility breaks are found in aggregate inflation for all 13 countries, with coefficient breaks in 12, 12 and

[14] All results discussed in this section employ the lags shown in this table.

[15] Hall and Sakkas (2013) provide formulae to approximate the asymptotic distributions which they show to deliver more accurate critical values than those provided by Bai and Perron (2003b) and have the additional advantage of allowing computation of asymptotic p-values.

10 countries for core, energy and food inflation, respectively. Although a maximum of five breaks is allowed, the most uncovered in any model is three coefficient breaks for each of aggregate, core and energy inflation in the US and for energy in Germany. Broadly similar numbers of volatility as coefficient breaks are uncovered, with three again the maximum number detected (for aggregate and energy inflation in Japan). Our finding of fewer breaks overall than univariate studies that examine the mean level of inflation (such as Benati, 2008, and Bataa *et al.*, 2014) may be attributed to our methodology[16] that explicitly includes dynamics in the breaks analysis in order to avoid the oversizing of HAC methods, as discussed in Section 3 above.

In order to establish the robustness of the above results we have also applied the iterative structural break testing algorithm to different definitions of the respective foreign inflation series as discussed in Section 2.1. Qualitatively the results of the structural break identification strategy remain unchanged. Most changes that arise are slight changes in the identified break dates. In Figure 3 we show the changes in identified break dates that arise for the US aggregate inflation series. When comparing the results for the bilateral trade weighted foreign inflation series (Panel a) and the results for the total trade weighted series (Panel b) we see that it is the break date for the last of three identified structural breaks that differs. The remaining two break dates remain essentially unchanged. When using the GDP weighted foreign inflation series (Panel c) the algorithm only identifies two of the three breaks.

As it turns out, in the illustrated case of the US (and the cases of other countries), even though we find different numbers of structural breaks, this does not change our conclusions about the general changes in the range of inflation characteristics and dynamics which will be discussed in the following Sections. In Figure 3 this can be seen by comparing the resulting, regime specific unconditional mean characteristics (*UcM*) and contemporaneous coefficients for the foreign inflation (*Contemp*, $\beta_{0j,s}$ in equation (6)). They display very similar changes across time regardless of whether the algorithm detects two or three regime changes.

4.2. Shortrun movements

As explained in Section 3, the globalisation of inflation has a number of implications for the nature of changes in the parameters of the model in equation (6) and this subsection considers our results in terms of the short-run changes in the coefficients of foreign inflation.

Figures 4 and 5 plot the estimated contemporaneous and (summed) lag foreign inflation coefficients, respectively, over regimes for the 13 countries of our sample. That is, for $\widehat{m}$ coefficient breaks detected in equation (6) for each specific country s, Figure 4 shows the value of $\widehat{\beta}_{0j}$ for each implied regime $j = 1, ..., \widehat{m} + 1$. The vertical axis measures the estimated coefficient value, while the horizontal axis represents time. Therefore, a detected coefficient break date (given in Table A.2) is indicated by a vertical shift in the estimated coefficient value. In addition to the graphical representation of Figure 4, the coefficient values are provided in Appendix Table A.5. Panels a, b, c and d of Figure 4 correspond to aggregate, core, energy and food inflation, respectively[17].

[16] An analysis of univariate aggregate monthly inflation for the countries used in this study and employing the same methodology finds similar numbers of breaks to those reported in Appendix Table A.2; detailed results are available from the authors on request.

[17] Equivalents to Tables A.1 and A.5 but using the alternative definitions of the foreign inflation variable can be found in the Web Appendix. The differences that arise from using these alternative definitions are not qualitatively different to those discussed here. This also applies for the remaining discussion.

Figure 5 is analogous to Figure 4, except that the information presented is the sum of the estimated lagged foreign inflation coefficients $\sum_{i=1}^{r} \widehat{\beta}_{ij}$ in equation (6) over regimes $j = 1, ..., \widehat{m} + 1$; the values plotted are also shown in Table A.5. Note, however, that lagged foreign inflation is not included in all models (Table 1), and hence fewer than 13 countries appear in the panels of Figure 5.

Panel a of Figure 4 supports the globalisation hypothesis for aggregate inflation, in the sense that the contemporaneous foreign inflation coefficient is almost always higher in the final coefficient regime compared with the 1970s for almost all countries under study[18]. Although the magnitudes and dates of change vary over countries, the overall pattern of positive and increasing comovement is clear. Therefore, our model reproduces the pattern of inflation globalisation documented in other studies that use aggregate CPI inflation, including Ciccarelli and Mojon (2010), Neely and Rapach (2011) and Bataa *et al.* (2014). It should be noted that this phenomenon apparently applies only in the very shortrun, since the summed coefficients of Figure 5 panel a indicate that shifts to lagged foreign inflation generally play a smaller, rather than greater, role over time.

Scrutiny of core inflation in panel b of Figure 4, however, does not lead to corresponding conclusions about globalisation. In contrast to the increases seen in panel a, the contemporaneous foreign coefficient is relatively constant over time for core inflation and of smaller magnitude than for the aggregate series. Indeed, with the exception of Canada, decreases in the contemporaneous foreign coefficient occur in all economies that introduced inflation targeting in the early 1990s (UK, Finland and to a smaller degree Sweden). Further, the six Euro area countries present no substantial evidence that foreign core inflation plays a greater role with monetary integration, except for a marked increase in the contemporaneous coefficient for Italy from 2004. Overall, the coefficients of lagged foreign core inflation in Figure 5 (panel b) similarly show little evidence of increase.

The implication is that non-core inflation elements must largely drive the increased contemporaneous comovement seen in aggregate inflation. Indeed, panels c and d of Figure 4 show similar characteristics in this respect to panel a, with increased comovement particularly clear for energy inflation (panel c). The large estimated contemporaneous foreign energy inflation coefficient for the US from 1993 is notable, with this being numerically very similar to the corresponding coefficients for the small countries of Finland and Switzerland in the latter part of the sample. With the exceptions only of Japan and Sweden, it is striking that the contemporaneous foreign energy inflation coefficient is higher at the end of the sample than at the beginning (in the 1970s) for every country we study. Although less clearcut, there is also an overall pattern of increase in the contemporaneous foreign coefficient in the food inflation model. Interestingly, both France and Germany experience an increased foreign food inflation role during the 1980s, and this may be due to greater monetary integration causing price movements from other member states to be transmitted more fully to these large countries.

4.3. Mean inflation

Shortrun movements are only one implication of inflation globalisation. To examine the longrun, Figure 6 shows mean inflation levels and how these change over time with coefficient regimes. More specifically, the values graphed are computed as the sample means of domestic (aggregate, core, energy

[18] As seen from Table A.5, Japan and the Netherlands are the only exceptions to this statement.

or food) inflation series over the regimes $j = 1, ..., \widehat{m} + 1$ detected in equation (6); the values plotted are included in Table A.5.

Panel a of Figure 6 presents strong evidence that all 13 OECD countries have effectively converged to a common aggregate inflation level of about 0.2 percent a month. Further, this common inflation level applies also to core inflation (panel b) in the latter part of the sample. Indeed, it is striking that each break in both panels is associated with a lowering of (aggregate or core) inflation in all countries. In addition to the general decline of the 1980s, Figures 1 and 6 (Panels a and b), indicate that all the countries in our sample that introduced inflation targeting in the early 1990s, (UK, Canada, Sweden and Finland), experienced a decline in the aggregate or core inflation in that period, (see also Table A2). For Italy, the change dated in May 1996 brings that country's mean inflation into line with the requirements of the Maastricht Treaty[19]. This finding emphasises not only the decline in domestic mean inflation previously documented in international inflation studies, such as Cecchetti and Debelle (2006), but also the communality of this new level.

The mean level of energy inflation (panel c of Figure 6) also often declines from its initially high level of the 1970s, while the level of food inflation is either constant over time or declines. The panels of Figure 6, therefore, show declines in the mean level of inflation to be general across the core, energy and food elements of the CPI in these countries. However, there appears to be more cross-country divergence in mean levels of energy inflation at the end of the sample than for other components.

Consequently, mean levels of aggregate inflation are compatible with the level of inflation being essentially a global phenomenon from around the mid-1990s, in contrast to the differing levels of the 1970s. This long-run aspect may explain why Ciccarelli and Mojon (2010) find that inclusion of a longrun error-correction to foreign inflation improves the accuracy of domestic forecasting models. In addition to the aggregate, the movement over time towards a cross-country alignment of mean core inflation is particularly striking, especially since shortrun fluctuations in core inflation do not show evidence of an increased role for foreign inflation.

Previous authors, including Neely and Rapach (2011) and Mumtaz and Surico (2012), suggest that monetary policy plays a role in explaining the globalisation of inflation, and our results support this in terms of the level of inflation. In particular, although the dates of structural change are estimated in terms of the domestic-foreign relationship of equation (6), those for core inflation at the beginning of the 1990s may be associated with the introduction of inflation targeting for each of Canada, Sweden, Finland and the UK. Although monetary policy traditionally focuses on underlying, or core inflation, most inflation targeting economies target directly aggregate CPI inflation. It is worth noting that with the exception of the UK, which targeted the RPIX until 2003, all other inflation targeters experienced structural changes in both core and aggregate CPI inflation in the early 1990s, when inflation targets were first introduced.

Whatever the immediate causes of a reduction in the level of inflation, inflation targeting may help to keep inflation expectations at a lower level and consequently make a downward shift permanent. There is, however, relatively little evidence that the introduction of the Euro altered mean levels of core inflation in member countries, except perhaps for the decline in the levels of aggregate inflation in

[19] The formal Maastricht Treaty requirement is for annual inflation to be no more than 1.5 percent higher than the average of the three lowest inflation countries of the European Union.

France and Italy around the mid-1990s. This is not surprising, since the common currency is feasible due to the earlier convergence in mean inflation rates across most member countries, rather than convergence being a consequence of the Euro currency.

4.4. Persistence and volatility

Changes in the characteristics of inflation persistence and (disturbance) volatility are depicted in Figures 7 and 8, respectively. In common with earlier figures, panels a to d show regime-specific values relating to each of aggregate, core, energy and food inflation. Regimes applying for persistence are again those implied by the coefficient break dates, while distinct regimes are permitted for volatility (given in Appendix Table A.2).

In line with results for international analyses that employ univariate inflation models (including Benati, 2008, and Bataa *et al.*, 2014), our results measured through the domestic-foreign inflation model in equation (6) indicate that aggregate inflation persistence has generally declined over this period. Note, however, that our measure of persistence is given by $\sum_{i=1}^{p} \widehat{\alpha}_{ij}$ and hence is conditional on foreign inflation. Panel a of Figure 7 shows every country (except Austria, where no domestic lags are selected) to share this decline, with persistence in aggregate inflation being 0.4 or less for all countries by the end of the sample. However, as for comovement in Figure 4, core inflation shows a different pattern for persistence in panel b of Figure 7 from the aggregate in panel a. More specifically, persistence for core inflation is often constant or increases and is also often higher for core than aggregate inflation in the respective final coefficient regimes. All inflation targeting economies, with the exception of the UK, that targeted the RPIX until 2003, experienced a reduction in aggregate inflation persistence in the early 1990s, whereas in the UK this reduction is captured in core inflation. Moreover, close inspection of Figure 7 b, indicates that all inflation targeters (except Finland) also experience a fall in core inflation persistence during the early 1990s. This is not true for most of the other Euro area economies where core inflation persistence appears to remain constant (with the exception of Italy). Thus, while a reduction in persistence is evident for aggregate inflation, there is no evidence of a similar reduction for core inflation, apart from the inflation targeting economies in the early 1990s

Further, our models find little change over time in persistence for energy inflation, with this often being zero as no lagged dependent variable is required (Table 1). Similarly, the model for food inflation requires no lags in a majority of cases, while persistence declines in others. These findings support those of Altissimo *et al.* (2006), who note less persistence for non-processed foods and energy and higher persistence for industrial goods and services in the Euro area, while Nason (2006) finds higher persistence for core than aggregate inflation in the US.

Volatility is an important, but often overlooked, feature of inflation. The disturbance volatility implications for the model of equation (6) are plotted in Figure 8 for each of the four series we analyse, but note the different scales used here for energy and food inflation versus aggregate and core. The volatility reductions detected around the early to mid-1980s are a manifestation of the international dimension of the so-called Great Moderation, which a number of studies link to improved monetary policy; see, for example, the discussion in Summers (2005). Such reductions apply particularly to aggregate and core inflation, with all countries experiencing volatility declines for at least one (and typically both) of these. This relatively constant cross-country volatility, which again contrasts with the 1970s, is compatible with globalisation of inflation. In general, energy inflation is volatile until the mid 1980s, with a somewhat mixed picture of declines and increases subsequently, with little evidence of convergence in volatility across countries. On the other hand, the volatility of food inflation typically declines, and is generally relatively stable from the mid-1990s.

The results of this section, therefore, support the proposition that the adoption of similar monetary policies across countries plays an important role in the globalisation of inflation, by effectively bringing the level and volatility of core inflation into line across countries. This would appear to be more true since the 1990s, rather than the earlier periods.

5. Decomposing Shortrun Comovements

Further insight into changes in international inflation linkages are provided by the model in equation (9). More specifically, this focuses on CPI inflation at the aggregate level, since the globalisation or comovement of inflation is documented using such data in the studies of Ciccarelli and Mojon (2010), Neely and Rapach (2011), Mumtaz and Surico (2012), Eichmeier and Pijenburg (2013), Bataa, Osborn, Sensier and van Dijk (2013), and others. However, to shed light on the nature of this comovement and how it changes over time, (country-specific) foreign inflation is decomposed into core, energy and food components.

The results discussed in the previous section suggest that the globalisation of inflation as measured by increased shortrun comovement in aggregate inflation is driven largely by the energy component of inflation and, to a lesser extent, the food component, with core inflation playing relatively little role. The decomposition provided by the model in equation (9) is designed to examine this further.

As discussed in Section 3 above, two sets of break dates are employed in this analysis, namely those obtained from the model in equation (6) applied to aggregate inflation and breaks estimated from application of our iterative procedure directly to the model in equation (9); the latter break dates (obtained using a 5% significance level) are included in Appendix Table A.2. Irrespective of which method is used to estimate the break dates, the dynamics for equation (9) are specified in the context of that model, allowing different lags to apply to each individual foreign inflation component series. These lags are included as the final columns of Table 1.

Consider first the lags selected by our SIC procedure (allowing a maximum of 12 lags for both the lagged domestic inflation variable and each of the foreign components) in the context of equation (9). With the contemporaneous values treated in the same way as any individual lag, current foreign core inflation is selected in only three of the 13 cases, which relate to the Euro area countries of Austria, Finland and France. This again supports the hypothesis that core inflation does not drive the increased shortrun international comovement of inflation. Nevertheless, lagged foreign core inflation is typically selected, although neither contemporaneous nor lagged foreign core inflation would be included in the models for Germany, Japan, Netherlands or Switzerland. The situation for foreign energy inflation stands in contrast, with the contemporaneous value always selected and a lagged value only for Germany. Finally, contemporaneous foreign food inflation is selected for a majority of countries (8 of 13), with no contemporaneous or lagged value selected for the UK. However, to facilitate comparison with the results discussed above for the model in equation (6), contemporaneous values of all foreign component series are included in the specification of equation (9) when this is employed for structural break testing. Although there are some differences, the broad patterns of domestic lags selected for equation (9) are similar to those for equation (6) estimated using aggregate CPI inflation.

Table 2 provides the results of the tests of coefficient equality, the methodology for which was discussed in subsection 3.4. To be specific, and conditional on the coefficient break dates used, Table 2 provides the estimated coefficients for foreign inflation in each coefficient regime. Also provided are p-values for tests of the null hypotheses of no contemporaneous coefficient break, for each component and jointly across all components, equations (10) and (11), together with the corresponding tests of equations (12) and (13) for lag coefficients. To facilitate comparison, the first line of results for each country in the table gives the coefficient estimates from the aggregate inflation model (as in Appendix Table A.5), together with the corresponding p-values for tests of coefficient equality across regimes applied to the aggregate inflation model in equation (6). As discussed in subsection 3.4, the estimated coefficients for foreign inflation components partly reflect the differing weights on components, so that those for energy and food are anticipated to be smaller than for core inflation, with that for the core reduced to a lesser extent, compared with those of the aggregate equation.

In the context of equation (6), eight countries have contemporaneous foreign aggregate inflation coefficients that exhibit significant change over regimes, all with very small p-values. Imposing the break dates from equation (6) on equation (9), the p-values for four of these countries (namely Austria, France, Germany and the US) imply that the disaggregate foreign component model attributes the change in contemporaneous effects to a changed role of foreign energy[20] and (to a lesser extent) foreign food inflation, and not to foreign core inflation. On the other hand, increases in the contemporaneous foreign inflation coefficients of the aggregate models in equation (6) for Italy, Sweden and Switzerland are associated by the disaggregated model in equation (9) with increases in the foreign core inflation coefficients, with no significant changes in the responses to foreign energy or food inflation; both core and energy coefficients exhibit significant change (both increases) for Canada. In no case does food inflation play a key role.

Many countries do not include lagged values of foreign inflation (aggregate or specific components) in the relevant models, and hence it is not surprising that there is there is much less evidence of change when lagged foreign coefficients are examined in the context of either model in equation (6) or equation (9). However, where change is significant in the context of the aggregate model (Austria, Canada, Italy), the foreign component coefficients generally decrease, implying a shift from lagged to contemporaneous linkages.

The second set of results in Table 2 are based on break dates estimated in the context of equation (9). Note that although the quoted p-values are asymptotically justified by the analysis of Bai and Perron (1998), they may over-state significance in a finite sample context, especially for a joint test applied to equation (9) when the break dates are also endogenously determined within this model; hence it is unsurprising to find greater apparent significance of change compared to when the break dates are imposed from equation (6). With this caveat, the results in Table 2 emphasise the important role of changes in the response to contemporaneous foreign energy inflation, with Denmark and the

[20]Our results on the increased role of energy for aggregate CPI inflation in the US contrasts with results of Hooker (2002), who finds a decreased role for oil prices after 1981. However, his context of a domestic Phillips curve model is quite different from the domestic-foreign linkages of our study.

UK being exceptions. Core inflation also plays a role for some European countries, namely Denmark, Italy, Sweden, Switzerland and the UK, although of these only Italy is a member of the Euro area. As measured by p-values, responses to foreign food inflation also sometimes play a role in changed linkages, namely for Finland, Sweden, the UK and US. Consideration of changes in lagged foreign coefficients does not substantially alter the pattern of results.

Results for changes in persistence and intercepts for equation (9) are not presented in order to conserve space[21]. However, these generally confirm the persistence decrease seen in the simpler model of equation (6) for aggregate inflation. When breaks are allowed in all the coefficients of equation (9), intercept changes are typically not significant.

There may appear to be some tension between the results for core inflation as presented in this section and the preceding one. More specifically, the results of Section 4 for core inflation do not provide evidence of globalisation in the shortrun movements of this component, whereas Table 2 shows significant changes in the coefficient of contemporaneous foreign core inflation for domestic aggregate inflation in some countries. Our interpretation is that, in focusing on shortrun movements, the model in equation (9) may incorrectly associate a convergence in mean levels of core inflation to changes in the contemporaneous coefficient of foreign core inflation. Therefore, the results of Table 2 serve to emphasise the roles of both core and energy inflation in understanding the apparent globalisation of inflation.

6. Conclusions

Previous analyses that document high international inflation linkages since the 1980s, including Ciccarelli and Mojon (2010), Neely and Rapach (2011), Bataa, Osborn, Sensier and van Dijk (2013), primarily focus on aggregate inflation. The present paper sheds new light on the nature of this so-called globalisation of inflation by examining the separate roles of core, energy and food inflation in the increased synchronicity of monthly CPI inflation across OECD countries. To do so, we analyse changes in the linkages of domestic and country-specific foreign inflation (the latter constructed as a weighted average) using an iterative methodology that allows for breaks in both coefficients and disturbance variances.

The principal results of our analysis can be summarised as follows:

- The longrun linkages between domestic and foreign CPI inflation have changed since the beginning of the 1970s, with the mean levels of both headline (aggregate) and core inflation effectively converging across countries from the early 1990s.
- The components of energy and food inflation largely drive shortrun (month-to-month) movements in aggregate inflation, while the comovement of core inflation is less marked than that of the aggregate series.
- The role of foreign energy inflation for explaining shortrun movements in domestic aggregate inflation has increased for many of the OECD countries we study, including the US.

[21] These may be obtained from the authors on request.

- Inflation volatility (represented as the variance of the residuals in our domestic-foreign inflation model) has declined over time for aggregate, core and food inflation. In contrast to the 1970s, volatility is relatively constant across countries from the 1990s for each of these three series.

- There is little evidence of convergence across countries or reduction over time in the volatility of energy inflation. Hence in a low inflation world, this component has become relatively more important for explaining shortrun movements in aggregate inflation.

These results, particularly the convergence in the mean level and volatility of inflation, but also the decline in inflation persistence since the early 1990s, suggest an important role has been played by monetary policies that focus on inflation, whether considered as headline or core inflation (Bullard, 2011). This link appears to be much stronger in the inflation targeting countries (Canada, Sweden, Finland and the UK), rather than in the Euro area, (with the exception perhaps of Italy), or in other economies. Overall, however, the apparent globalisation of inflation may be a consequence of the adoption of similar monetary policies across countries, rather than increased transmission of foreign inflation to individual countries.

The increased role identified for foreign energy inflation in explaining shortrun movements in domestic aggregate CPI inflation is a key result of our study. Bernanke (2007) suggests that the globalisation of inflation requires policymakers to pay increased attention to movements in foreign inflation, and our results imply that international energy price inflation should be the particular focus of this attention.

In summary, by studying the major components of core, energy and food inflation, our analysis of structural breaks in international inflation linkages reveals that the globalisation of aggregate CPI inflation has two key drivers. Firstly, there has been a convergence in the mean level of core (that is, excluding energy and food) inflation. To be specific, an effectively common level of the longrun mean (and also volatility) for core inflation has applied across the OECD countries of our sample since the late 1990s. Secondly, increased shortrun comovement seen since the 1980s is driven largely by individual countries responding more strongly, and quickly, to foreign movements in energy inflation.

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Tables

Table 1. Lags selected in domestic/foreign inflation models

Country	Aggregate Inflation		Core Inflation		Energy Inflation		Food inflation		Domestic Aggregate-Foreign Components			
	Domestic	Foreign	Domestic	Foreign	Domestic	Foreign	Domestic	Foreign	Domestic	Foreign Core	Foreign Energy	Foreign Food
Austria	NA	0,4	1,6,10	0*,2	11	0,1	NA	0	7	0,4	0	0
Canada	5,7,9	0,9	2,3,11	0*,1	10	0,4	9	0,1	5,7	0*,1	0	0
Denmark	1,6,11	0	1,3,6,10	0	NA	0,2	NA	0,1	1,6	0*,10	0	0
Finland	2,9	0,4	3,7	0,5	NA	0	4	0,4	2,9,11	0,2	0	0*,4
France	1,3,8	0	1,2,9	0,5	1	0	4,9,10	0	1,3,8,10	0	0	0
Germany	6	0	2,3,8	0*,1	NA	0	NA	0	2,3,6	0*	0,9	0
Italy	1,3,6	0,2	1,2,3,9	0*,2	3	0,1	1,6,7	0,2	1,3,6	0*,2	0	0
Japan	3,5,9,11	0,1	3,4,5,6,10	0*	1	0,1	9,11	0,11	1,5,9,11	0*	0	0*,1
Netherlands	4,6,8	0	4,5,6,9	0	NA	0,10	10	0	4,5,6,7,8	0*	0	0
Sweden	7,8,9	0	2,5,8,11	0*,2	NA	0	2,8,9	0	7,8,9	0*,5	0	0*,1
Switzerland	1,6,9	0	1,2,6,9	0*	NA	0	7	0	1,2,6,9,10	0*	0	0*,7
UK	1,2,3	0	1,2,5	0,5	1,8	0	1,7	0,5	1,2,6	0*,7	0	0*
US	1	0	1,2,8	0*,4	1,11	0	1,9	0	1,6	0*1	0	0

Notes: The aggregate, core, energy and food inflation models refer to equation (6), while the domestic aggregate with foreign components model is given by equation (9). Individual lags are selected to a maximum of $p = r = 12$. NA indicates that no lagged domestic values are selected. * indicates the contemporaneous value (lag 0) of foreign inflation is not selected, but is included in the estimated model. All estimations employ percentage monthly inflation data series.

Table 2. Estimates of foreign inflation coefficients in aggregate and aggregate-foreign components models

Country	Coefficients	Aggregate inflation model break dates												Aggregate-foreign components model break dates											
		Contemporaneous coefficients				Equality p-value	Equality joint p-value	Lag coefficients				Equality p-value	Equality joint p-value	Contemporaneous coefficients				Equality p-value	Equality joint p-value	Lag coefficients				Equality p-value	Equality joint p-value
		R1	R2	R3	R4			R1	R2	R3	R4			R1	R2	R3	R4			R1	R2	R3	R4		
Austria	Aggregate model	0.42	0.74			0.00		0.30	0.01			0.01		NA	NA			NA		NA	NA			NA	
	Foreign core	0.14	0.22			0.63		0.28	0.00			0.09		0.19	0.17			0.89		0.26	-0.02			0.06	
	Foreign energy	0.01	0.07			0.00	0.00	0.02	0.01			0.64	0.22	0.01	0.07			0.00	0.00	NA	NA			NA	0.06
	Foreign food	0.26	0.11			0.03		NA	NA			NA		0.22	0.12			0.08		NA	NA			NA	
Canada	Aggregate model	0.23	0.68			0.00		0.29	0.07			0.03		NA	NA			NA		NA	NA			NA	
	Foreign core	-0.09	0.52			0.00		0.36	0.07			0.17		-0.06	0.01			0.68		0.50	0.21			0.12	
	Foreign energy	0.00	0.07			0.00	0.00	0.03	0.00			0.06	0.01	0.00	0.07			0.00	0.00	NA	NA			NA	0.12
	Foreign food	0.13	0.21			0.11		0.05	-0.07			0.02		0.12	0.13			0.75		NA	NA			NA	
Denmark	Aggregate model	0.37	0.47			0.66		NA	NA			NA		NA	NA			NA		NA	NA			NA	
	Foreign core	0.21	0.01			0.50		NA	NA			NA		0.56	-0.13			0.00		-0.04	0.02			0.66	
	Foreign energy	0.02	0.07			0.35	0.66	0.09	0.00			0.05	0.13	0.08	0.07			0.87	0.00	NA	NA			NA	0.66
	Foreign food	-0.05	0.12			0.34		0.08	-0.01			0.55		0.11	0.10			0.94		NA	NA			NA	
Finland	Aggregate model	0.60	0.71			0.31		0.17	0.12			0.56		NA	NA	NA		NA		NA	NA	NA		NA	
	Foreign core	0.47	0.28			0.35		0.31	0.00			0.08		0.61	0.48	0.24		0.34		0.18	0.41	0.08		0.36	
	Foreign energy	0.05	0.08			0.26	0.17	NA	NA			NA	0.21	0.15	0.02	0.08		0.03	0.01	NA	NA	NA		NA	0.02
	Foreign food	-0.04	0.16			0.05		0.02	0.07			0.60		-0.30	0.01	0.17		0.01		0.21	-0.28	0.06		0.01	
France	Aggregate model	0.32	0.47	0.88		0.00		NA	NA	NA		NA		NA	NA			NA		NA	NA			NA	
	Foreign core	0.10	0.33	0.34		0.25		0.12	0.08	0.07		0.95		0.21	0.30			0.40		NA	NA			NA	
	Foreign energy	0.04	0.08	0.08		0.01	0.00	NA	NA	NA		NA	0.95	0.03	0.08			0.00	0.00	NA	NA			NA	NA
	Foreign food	0.09	0.16	0.17		0.42		NA	NA	NA		NA		0.08	0.18			0.07		NA	NA			NA	
Germany	Aggregate model	0.13	1.00	0.72		0.00		NA	NA	NA		NA		NA	NA	NA		NA		NA	NA	NA		NA	
	Foreign core	-0.11	-0.04	0.23		0.19		0.14	0.53	0.24		0.37		-0.05	0.21	0.25		0.11		NA	NA	NA		NA	
	Foreign energy	0.01	0.12	0.07		0.00	0.01	NA	NA	NA		NA	0.37	-0.01	0.13	0.07		0.00	0.00	0.00	0.02	0.02		0.59	0.59
	Foreign food	0.11	0.20	0.14		0.81		NA	NA	NA		NA		0.14	0.24	0.12		0.50		NA	NA	NA		NA	
Italy	Aggregate model	0.07	0.20	0.39		0.00		0.57	0.12	0.13		0.01		NA	NA	NA		NA		NA	NA	NA		NA	
	Foreign core	-0.42	-0.11	0.16		0.03		0.49	0.23	0.03		0.10		-0.20	0.22	0.21		0.02		0.44	0.24	-0.13		0.00	
	Foreign energy	0.03	0.03	0.03		1.00	0.27	0.02	0.01	0.02		0.85	0.10	0.04	0.01	0.05		0.00	0.00	NA	NA	NA		NA	0.00
	Foreign food	0.12	0.10	0.06		0.69		0.19	0.08	0.05		0.38		0.16	0.05	0.10		0.16		NA	NA	NA		NA	

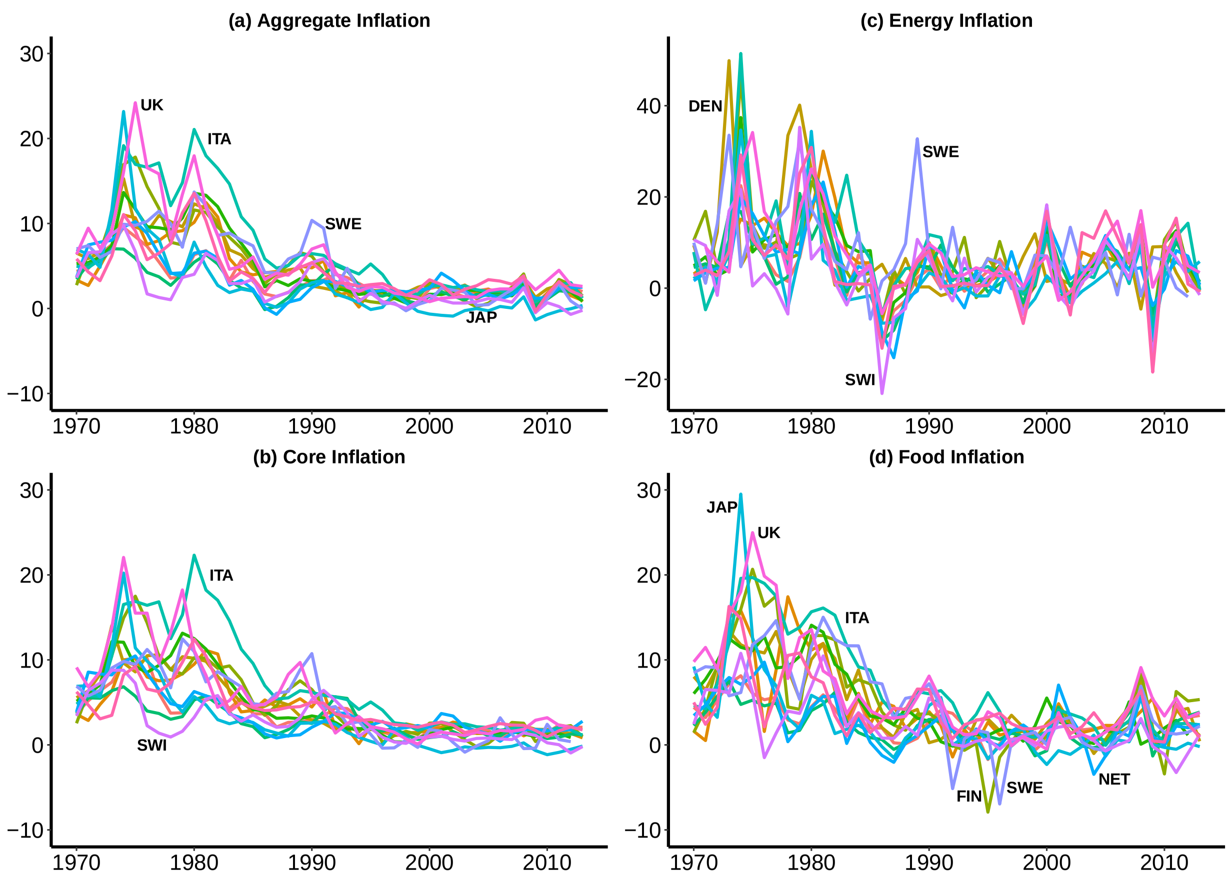
Table 2 continued

Country	Coefficients	Estimated break dates from aggregate inflation model												Estimated break dates from aggregate-foreign components model											
		Contemporaneous coefficients				Equality p-value	Equality joint p-value	Lag coefficients				Equality p-value	Equality joint p-value	Contemporaneous coefficients				Equality p-value	Equality joint p-value	Lag coefficients				Equality p-value	Equality joint p-value
		R1	R2	R3	R4			R1	R2	R3	R4			R1	R2	R3	R4			R1	R2	R3	R4		
Japan	Aggregate model	0.31	0.29			0.54		0.63	0.21			0.21		NA	NA			NA		NA	NA			NA	
	Foreign core	-0.05	0.44			0.31		NA	NA			NA		-0.10	0.43			0.30		NA	NA			NA	
	Foreign energy	-0.04	0.02			0.62	0.29	0.46	0.01			0.00	0.00	0.28	0.03			0.01	0.06	NA	NA			NA	
	Foreign food	0.21	0.03			0.11		0.17	0.01			0.15		0.06	0.03			0.77		0.15	0.04			0.32	0.32
Netherlands	Aggregate model	0.45	0.39			0.80		NA	NA			NA		NA				NA		NA				NA	
	Foreign core	0.25	0.11			0.28		NA	NA			NA		0.09				NA		NA				NA	
	Foreign energy	0.04	0.05			0.58	0.53	0.00	0.01			0.48	0.48	0.05				NA	NA	NA				NA	NA
	Foreign food	0.16	0.12			0.58		NA	NA			NA		0.14				NA		NA				NA	
Sweden	Aggregate model	0.42	0.81			0.00		NA	NA			NA		NA	NA			NA		NA	NA			NA	
	Foreign core	0.17	0.95			0.02		0.12	0.44			0.27		-0.21	0.32			0.05		0.04	0.46			0.08	
	Foreign energy	0.08	0.08			0.88	0.01	NA	NA			NA	0.27	0.13	0.07			0.12	0.01	NA	NA			NA	0.04
	Foreign food	-0.06	0.15			0.16		NA	NA			NA		-0.11	0.16			0.04		0.26	0.02			0.05	
Switzerland	Aggregate model	0.28	0.87			0.00		NA	NA			NA		NA	NA			NA		NA	NA			NA	
	Foreign core	-0.38	0.44			0.00		NA	NA			NA		-0.43	0.05			0.01		NA	NA			NA	
	Foreign energy	0.13	0.09			0.17	0.00	NA	NA			NA	NA	0.13	0.09			0.02	0.01	NA	NA			NA	0.65
	Foreign food	0.26	0.14			0.27		NA	NA			NA		0.18	0.12			0.48		0.08	0.04			0.65	
UK	Aggregate model	0.56	0.59			0.46		NA	NA			NA		NA	NA	NA	NA	NA		NA	NA	NA	NA	NA	
	Foreign core	0.52	0.36			0.69		0.58	0.08			0.22		0.34	1.46	-0.11	0.20	0.01		0.96	-0.32	0.41	-0.02	0.00	
	Foreign energy	0.06	0.05			0.95	0.95	NA	NA			NA	0.26	-0.01	0.08	0.08	0.05	0.30	0.00	NA	NA	NA	NA	NA	0.00
	Foreign food	0.01	0.10			0.60		0.12	0.04			0.64		-0.04	-0.30	0.35	0.07	0.00		NA	NA	NA	NA	NA	
US	Aggregate model	0.63	0.46	0.37	1.36	0.00		NA	NA	NA	NA	NA		NA	NA	NA		NA		NA	NA	NA		NA	
	Foreign core	0.33	-0.01	0.09	-0.05	0.30		0.24	0.23	0.09	0.19	0.72		0.00	0.11	0.40		0.10		0.07	0.02	-0.07		0.71	
	Foreign energy	0.07	0.07	0.08	0.17	0.00	0.00	NA	NA	NA	NA	NA	0.72	0.07	0.09	0.15		0.00	0.00	NA	NA	NA		NA	0.71
	Foreign food	0.15	0.13	0.01	0.00	0.01		NA	NA	NA	NA	NA		0.14	0.00	0.08		0.01		NA	NA	NA		NA	

Notes: The aggregate model is equation (6) estimated using aggregate inflation, with other results referring to the aggregate-foreign components model in equation (9). The equality p-value refers to a test of the null hypothesis of equation (10) or (12) for contemporaneous and lagged foreign coefficients, respectively, while the equality joint p-value refers to the hypothesis of equation (11) or (13) tested over foreign core, energy and food coefficients. All p-values are obtained using the conventional F-distribution, conditional on estimated aggregate or aggregate-foreign component model break dates shown in Table A.2. The prefix R indicates regimes as implied by the respective break dates.

Figures

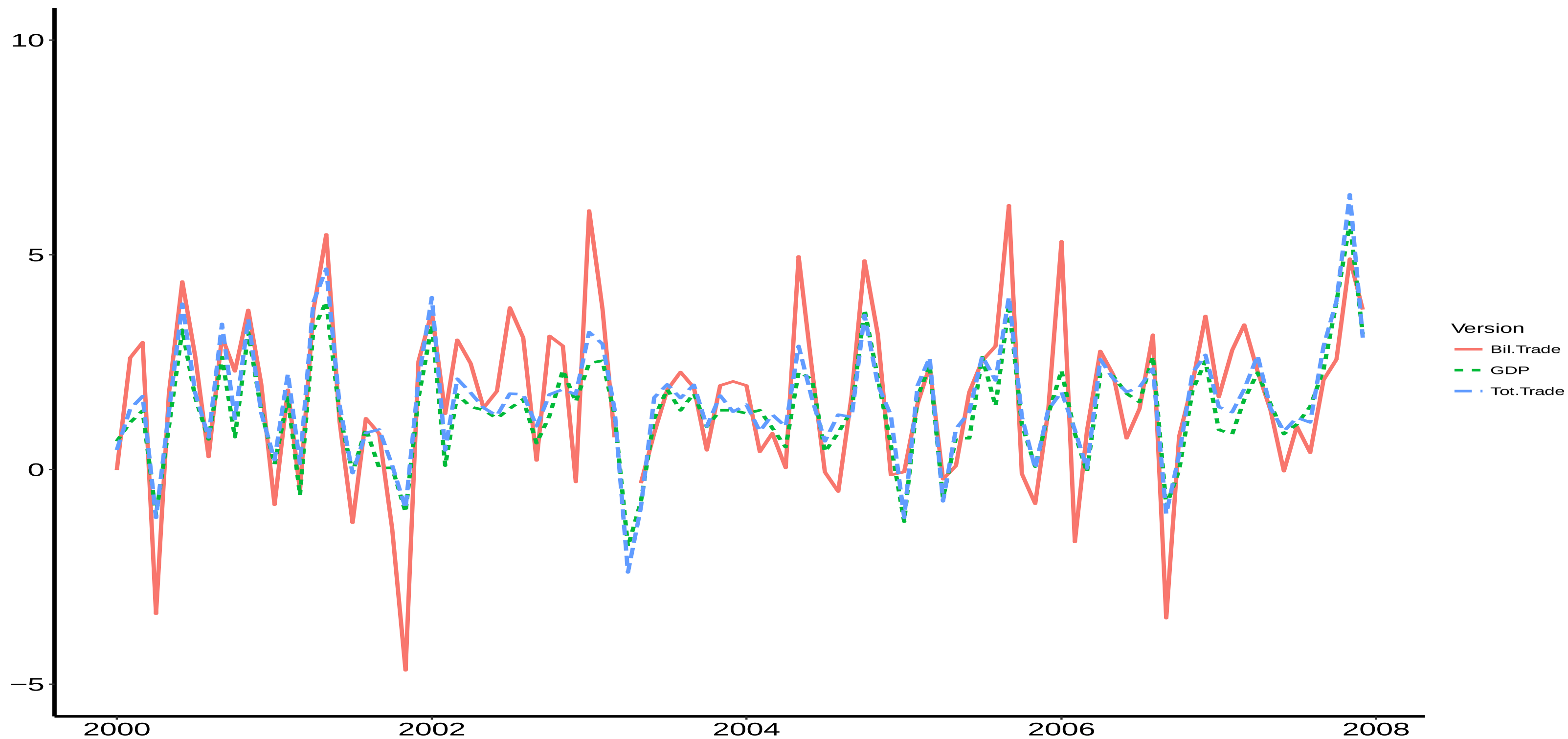
Figure 1: Domestic Inflation Dynamics



Notes: The values in the graph are annual percentage inflation (note the different scales for the Energy component in Panel (c)). Annual series are shown to remove higher frequency variation and make the Figure easier to read. The analysis in the paper uses monthly inflation series.

These figures are to display the general pattern of the respective inflation series in the 13 countries used in this paper. However, some series have been labelled with country codes: *DEN*: Denmark, *FIN*: Finland, *ITA*: Italy, *JAP*: Japan, *NET*: Netherlands, *SWE*: Sweden, *SWI*: Switzerland, *UK*: United Kingdom.

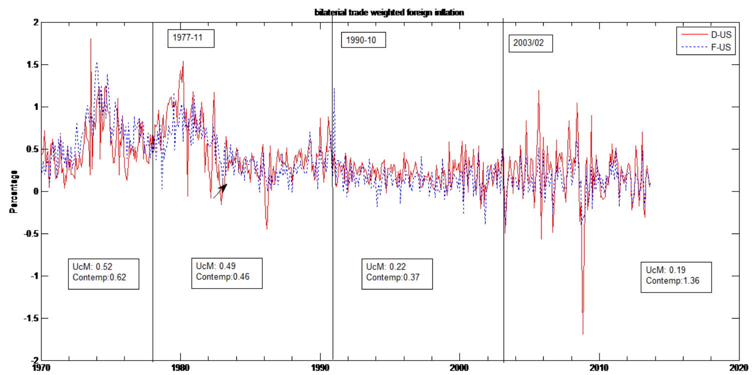
Figure 2: Foreign Inflation for the United States



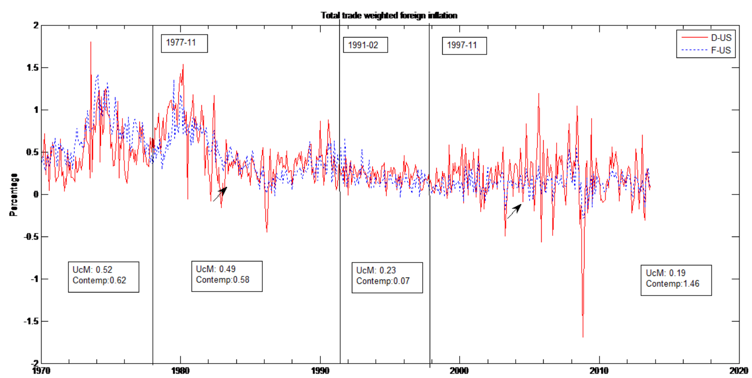
Notes: Foreign Inflation for the United States (US) using different definitions of the weighting variables. This Figure shows the subsample of monthly inflation (annualised) to make it easier to see the different characteristics. The differences in this subsample are representative of the differences across the entire sample.

Figure 3: Inflation Series and identified break dates.

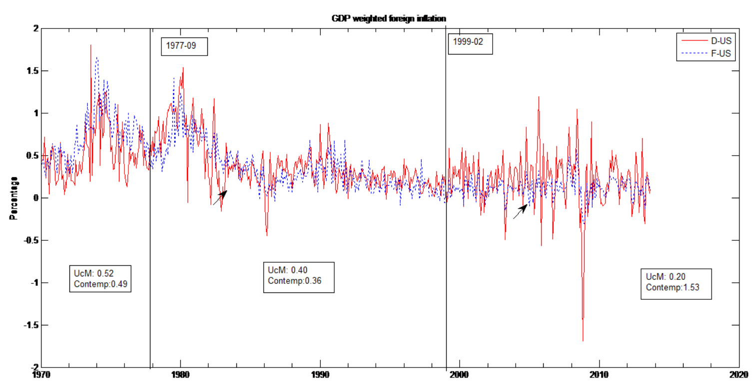
Panel (a)



Panel (b)

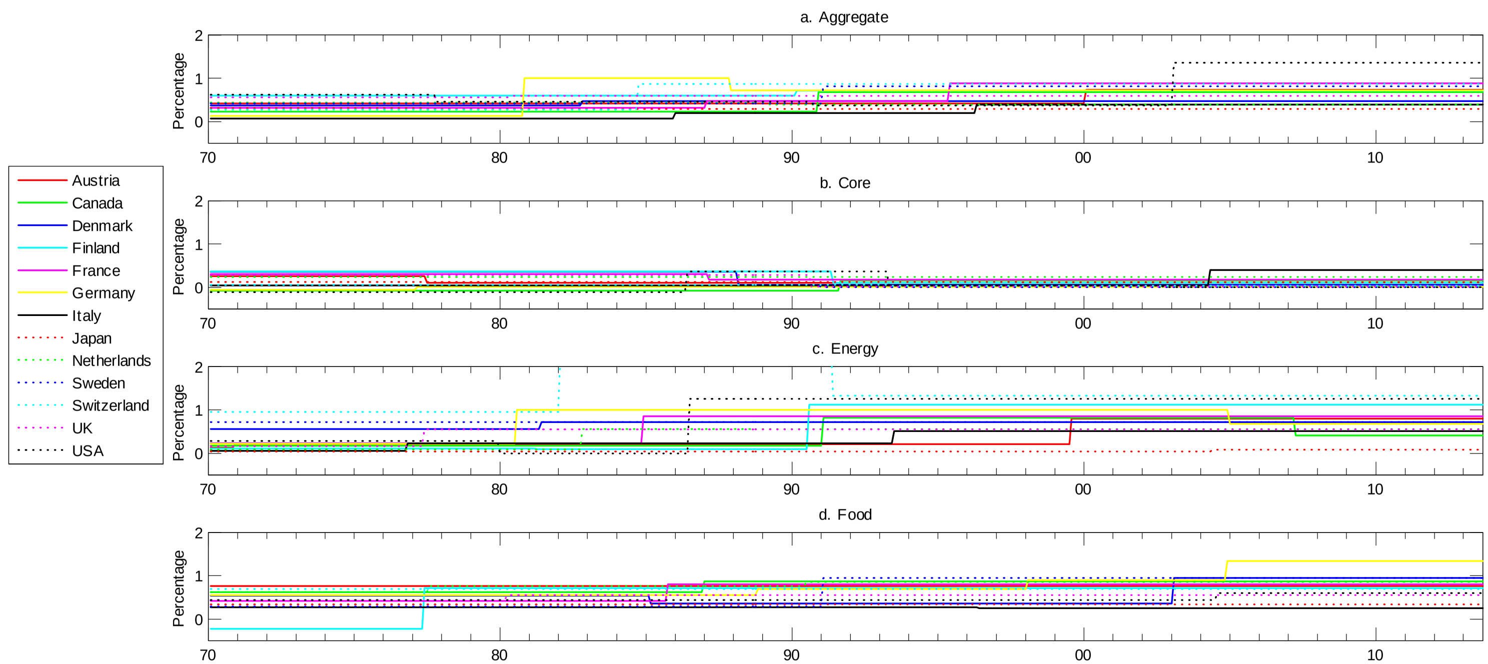


Panel (c)



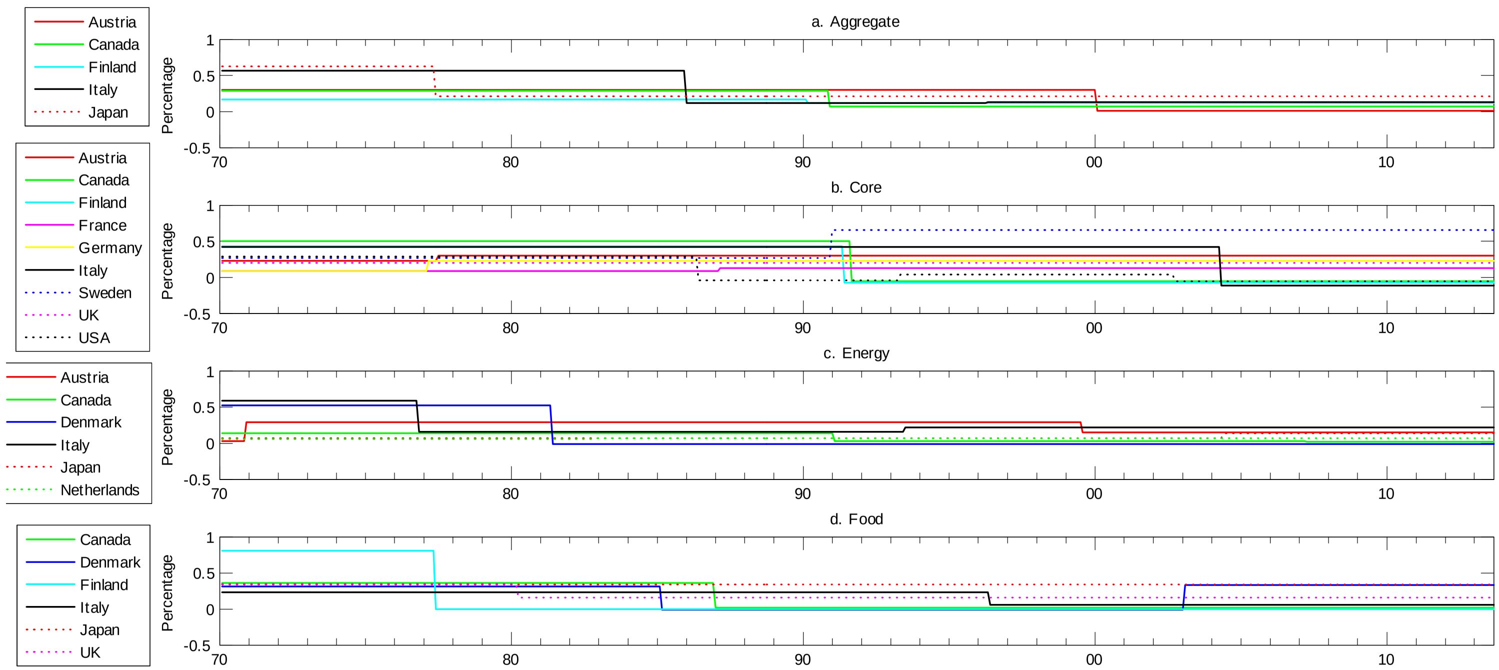
Notes: We display the domestic $(D-US)$ and foreign inflation $(F-US)$ for the US and the identified structural break dates. The three panels differ in the weights used to calculate $F-US$: (a): bilateral trade weights; (b): Total Trade weights; (c): GDP weights. Vertical lines represent identified structural breaks. *UcM* indicates the unconditional mean in an identified regime. *Contemp* identifies the estimate of the contemporaneous coefficient for foreign inflation in the respective regime.

Figure 4: Contemporaneous Foreign Inflation Coefficients



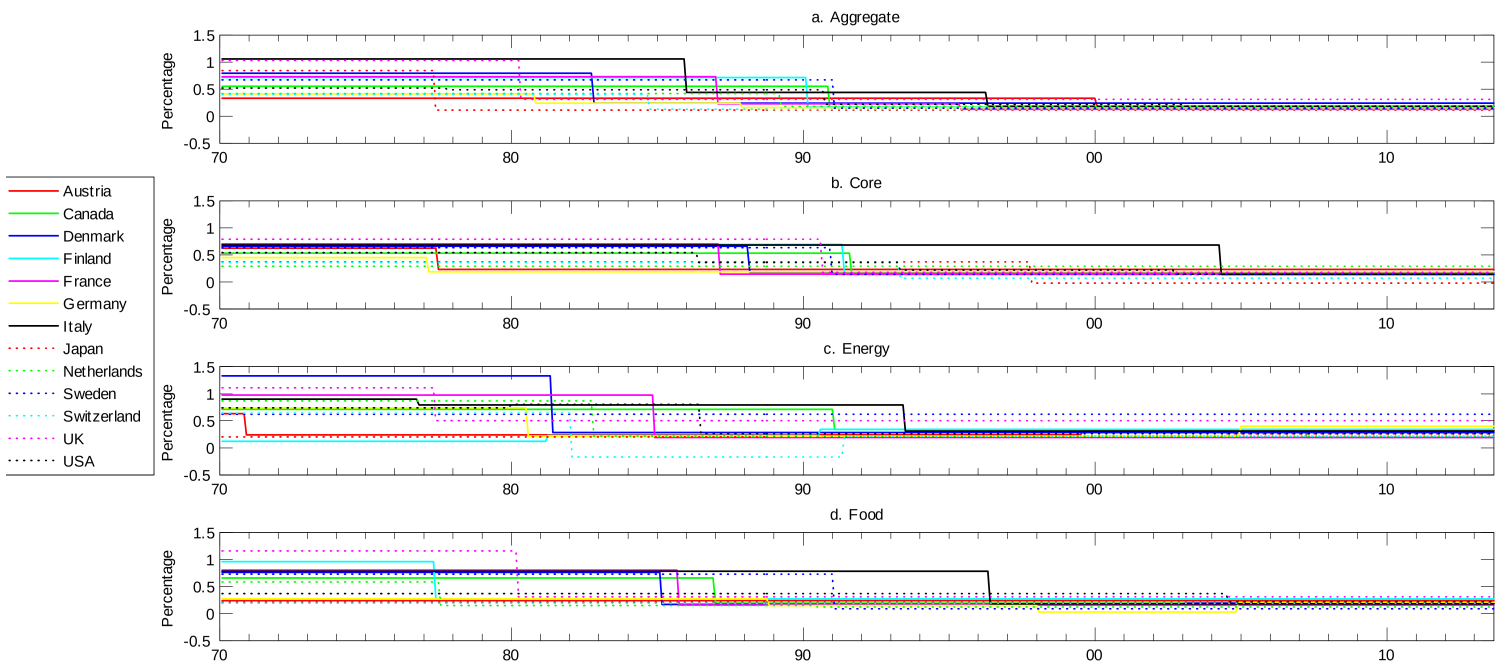
Notes: Coefficients are those of Table A.5, over regimes as defined by the break dates of Table A.2.

Figure 5: Lagged Foreign Inflation Coefficients



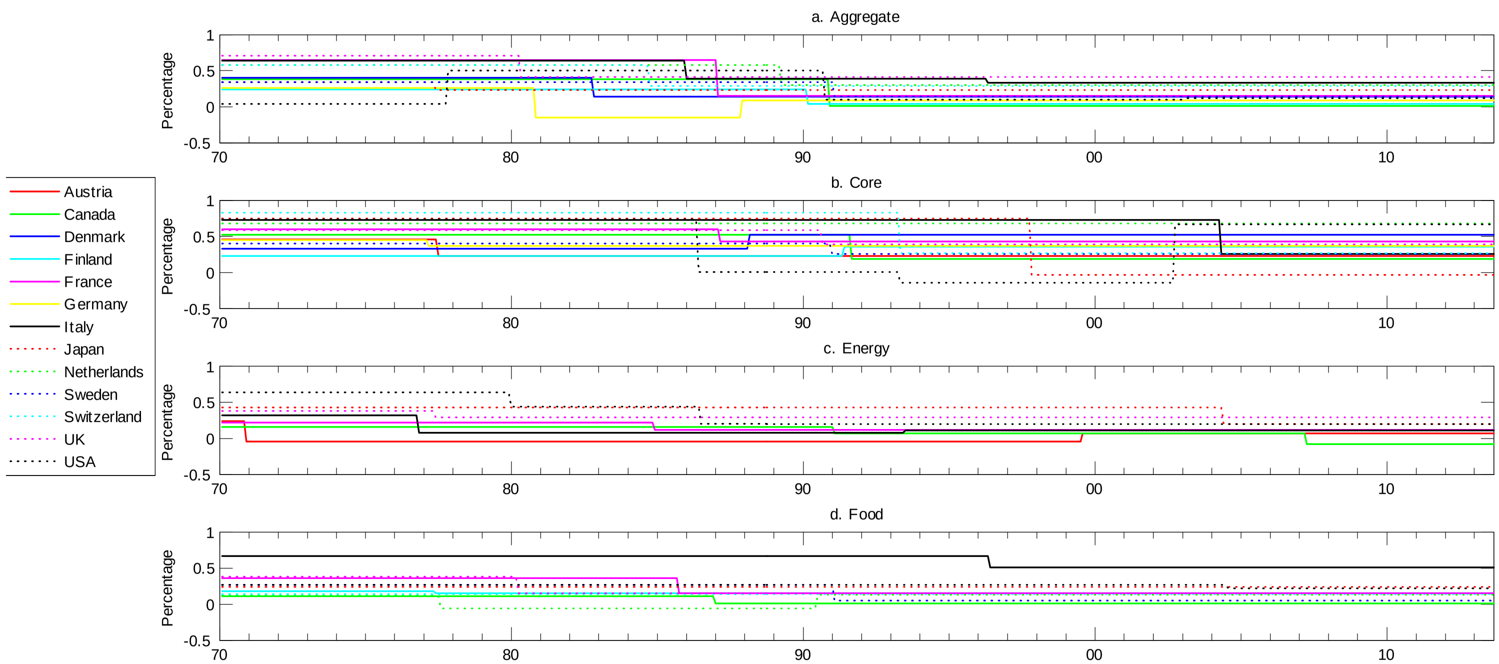
Notes: As for Figure 4.

Figure 6: Mean of Domestic Inflation



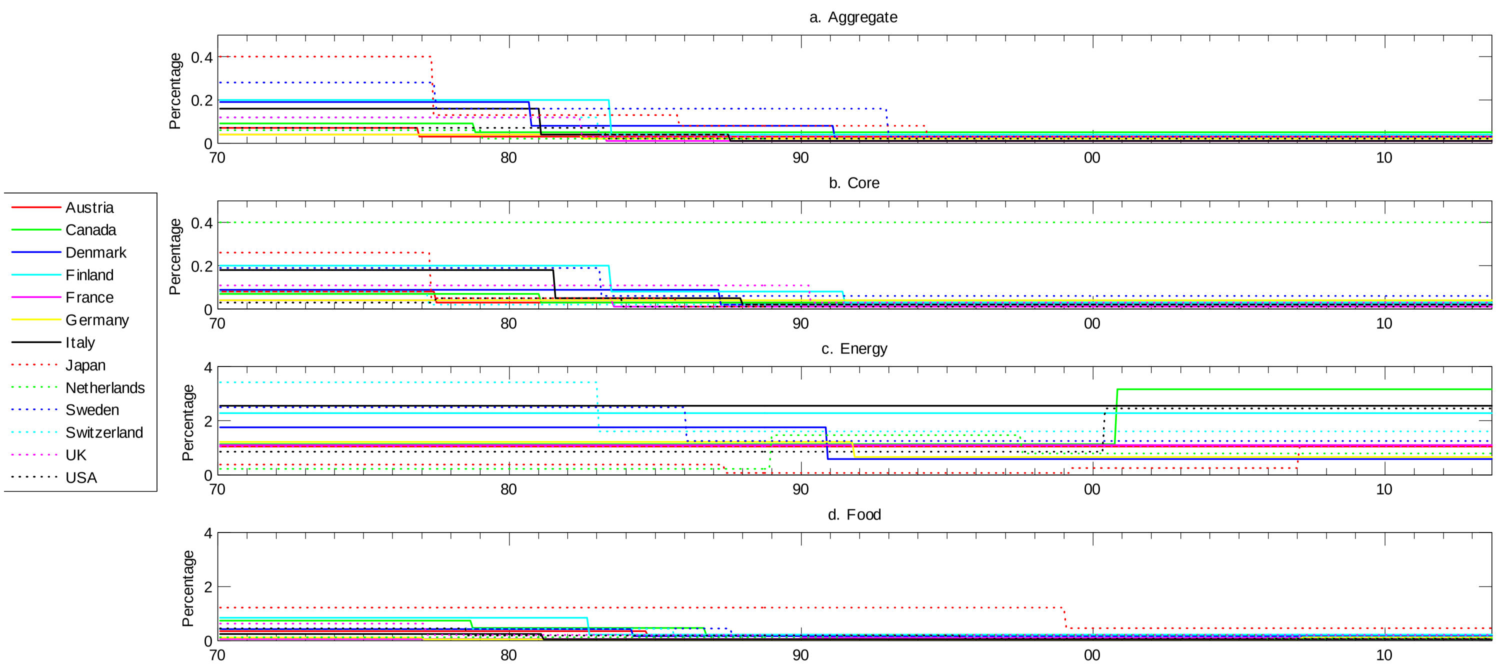
Notes: As for Figure 4.

Figure 7: Persistence



Notes: As for Figure 4.

Figure 8: Residual Variances



Notes: As for Figure 4.

Appendix

Detailed Algorithm

In detail, the algorithm of subsection 3.2 is as follows:

Step 1: Outlier detection. After the initial iteration, outlier identification for $\pi_{t,s}^{D}$ uses data scaled to account for volatility breaks (by dividing by the regime-specific standard deviation) and considers outliers separately for each subsample defined by the coefficient breaks, using break dates estimated for (6) in the previous iteration. Outliers are defined as four times the interquartile range from the median and are replaced by the median of the six neighboring non-outlier values. After outlier replacement, the series are rescaled to the original units. The initial iteration applies the Stock and Watson (2003) procedure over the whole sample without scaling.

Step 2: Model selection. The dynamics of the model (6) are specified using a general to specific multipath search algorithm (Hendry and Krolzig, 2001) in combination with the Schwartz Information Criterion (SIC). The intercept is always included and the model is estimated with $p = r = 12$. Five starting points are then generated by initially eliminating the single lag that is the i^{th} least significant ($i = 1, ..., 5$) in the general regression, calculating the corresponding SIC value in each case. From each starting point, the least significant variable is dropped sequentially one at a time, until only the intercept remains. The selected model is that which achieves the smallest SIC across all 25 models and 5 paths[22].

Step 3: Preliminary coefficient break test. After specifying the lags included in (6), the Bai and Perron (1998) multiple structural breaks test procedure is applied to the coefficients (intercept and all slope coefficients) employing heteroscedasticity consistent (HC) inference.

Step 4: Variance break test. Using the residuals from the model with coefficient breaks as identified in step 3, variance breaks are examined through tests applied to the mean of the squared residuals (see subsection 3.3).

Step 5: Coefficient break test. Since HC inference can lead to oversized coefficient break tests (Bai and Perron, 2006), breaks in the coefficients are reconsidered conditional on the variance breaks from step 4. Following the proposal of Pitakaris (2004), this is achieved by applying homoskedastic inference in (6) after applying the feasible GLS (generalized least squares) transformation. If no volatility breaks are detected, coefficient tests are applied to the original (outlier-corrected) data with a homoskedastic variance assumption.

[22]A comparison of this SIC based procedure with a conventional testing down method led to the selection of very similar lags.

Additional Tables

Table A1. Average bilateral trade weights by partner countries

	Aus	Can	Den	Fin	Fra	Ger	Ita	Jap	Neth	Swe	Swi	UK	US	SUM
Aus	0.000	0.008	0.012	0.010	0.056	0.525	0.118	0.026	0.038	0.026	0.078	0.050	0.054	1.000
Can	0.002	0.000	0.002	0.002	0.013	0.022	0.011	0.054	0.009	0.004	0.005	0.037	0.840	1.000
Den	0.015	0.009	0.000	0.038	0.066	0.280	0.054	0.035	0.074	0.183	0.023	0.142	0.080	1.000
Fin	0.017	0.012	0.050	0.000	0.067	0.218	0.049	0.045	0.068	0.207	0.026	0.145	0.095	1.000
Fra	0.015	0.014	0.013	0.009	0.000	0.314	0.175	0.037	0.090	0.025	0.055	0.128	0.125	1.000
Ger	0.076	0.013	0.030	0.016	0.187	0.000	0.129	0.045	0.153	0.039	0.073	0.110	0.129	1.000
Ita	0.039	0.015	0.014	0.008	0.224	0.304	0.000	0.027	0.074	0.020	0.066	0.091	0.117	1.000
Jap	0.006	0.067	0.009	0.006	0.039	0.097	0.028	0.000	0.036	0.013	0.024	0.062	0.614	1.000
Neth	0.015	0.009	0.021	0.014	0.136	0.411	0.072	0.028	0.000	0.032	0.023	0.137	0.102	1.000
Swe	0.019	0.014	0.115	0.091	0.077	0.230	0.051	0.037	0.079	0.000	0.027	0.149	0.112	1.000
Swi	0.054	0.011	0.013	0.009	0.136	0.348	0.123	0.044	0.049	0.023	0.000	0.082	0.108	1.000
UK	0.013	0.039	0.030	0.021	0.140	0.206	0.074	0.050	0.124	0.050	0.048	0.000	0.205	1.000
US	0.006	0.399	0.007	0.005	0.052	0.101	0.042	0.226	0.040	0.015	0.023	0.086	0.000	1.000

Note: The shares for each country with respect to partner countries are given in the rows of the table. The values shown are computed as monthly averages of trade shares over the sample January 1970-September 2013.

Table A.2. Estimated dates of structural breaks

Country	Model	Breaks in coefficients				Breaks in disturbance variance			
		1970s	1980s	1990s	2000s	1970s	1980s	1990s	2000s
Austria	Aggregate inflation	.	.	.	2000-Feb	1976-Dec	.	.	.
	Core inflation	1977-Jul	.	.	.	1977-Jul	.	.	.
	Energy inflation	1976-Apr	.	1999-Aug	.	.	.	.	.
	Food inflation	.	.	.	.	.	1984-Oct	.	.
	Aggregate-foreign components	.	.	1999-Oct	.	1977-Dec	.	.	.
Canada	Aggregate inflation	.	.	1990-Dec	.	1978-Nov	.	.	.
	Core inflation	.	.	1991-Sep	.	.	1981-Feb	.	.
	Energy inflation	.	.	1991-Feb	2007-Apr	.	.	.	2000-Nov
	Food inflation		1987-Jan	.	.	1978-Oct	1986-Oct	.	.
	Aggregate-foreign components	.	1989-Oct		.	.	1981-Feb	.	.
Denmark	Aggregate inflation	.	1982-Nov	.	.	.	1980-Oct	1990-Mar	.
	Core inflation	.	1988-Mar	.	.	.	1987-Apr	.	.
	Energy inflation	.	1981-Jun	.	.	.	.	1990-Dec	.
	Food inflation	.	1985-Mar	.	2003-Feb	.	1984-Apr	.	.
	Aggregate-foreign components	.	1989-Dec	.	.	.	1984-Apr	1990-Dec	.
Finland	Aggregate inflation	.	.	1990-Mar	.	.	1983-Jul	.	.
	Core inflation	.	.	1991-Jun	.	.	1983-Jul	1991-Jul	.
	Energy inflation	.	1981-Apr	1990-Aug	.	.	.	.	.
	Food inflation	1977-Jun	.	.	.	.	1982-Oct	.	.
	Aggregate-foreign components	1977-Sep	1989-Aug	.	.	.	1983-Jul	.	.
France	Aggregate inflation	.	1987-Feb	1995-Jun	.	.	1983-May	.	.
	Core inflation	.	1987-Mar	.	.	.	1983-Aug	1994-Nov	.
	Energy inflation	.	1984-Dec	.	.	.	.	.	.
	Food inflation	.	1985-Oct	.	.	.	.	1990-Jan	.
	Aggregate-foreign components	.	.	1990-Oct	.	.	1983-May	.	.
Germany	Aggregate inflation	.	1980-Nov 1987-Dec	.	.	.	1982-Jul	.	.
	Core inflation	1977-Mar	.	.	.	.	.	.	.
	Energy inflation	.	1980-Aug	.	2005-Jan	.	.	1991-Nov	.
	Food inflation	.	1988-Nov	1998-Feb	2004-Dec	1977-Feb	.	.	2007-Mar
	Aggregate-foreign components	.	1980-Feb 1988-Jan	.	.	.	1982-Jul	.	.
Italy	Aggregate inflation	.	1986-Jan	1996-May	.	.	1981-Feb 1987-Aug	.	.
	Core inflation	.	.	.	2004-May	.	1981-Aug 1987-Nov	.	.
	Energy inflation	1976-Nov	.	1993-Jul	.	.	.	.	.
	Food inflation	.	.	1996-Jun	.	.	1981-Mar	.	.
	Aggregate-foreign components	.	.	1996-Jul	2006-Oct	.	1981-Feb 1987-Oct	.	.
Japan	Aggregate inflation	1977-Jun	.	.	.	1977-Jun	1985-Nov	1994-May	.
	Core inflation	.	.	1997-Nov	.	1977-May	1985-Oct	.	.
	Energy inflation	.	.	.	2004-Jun	.	1987-Jun	1999-Apr	2007-Feb
	Food inflation	.	.	.	.	.	.	1999-Feb	.
	Aggregate-foreign components	1977-Jun	.	.	.	1977-Jun	1985-Dec	1994-May	.

Table A.2 continued

Country	Model	Breaks in coefficients				Breaks in variances			
		1970s	1980s	1990s	2000s	1970s	1980s	1990s	2000s
Netherlands	Aggregate inflation	.	1989-Apr	.	.	1979-Feb	.	.	.
	Core inflation	.	.	.	.	.	.	.	.
	Energy inflation	.	1982-Nov		.	.	1989-Jan	1997-Aug	.
	Food inflation	1977-Aug	.	1990-Jul	.	.	.	1991-May	2005-Dec
	Aggregate-foreign components	.	.	.	.	.	.	.	.
Sweden	Aggregate inflation	.	.	1991-Feb	.	1977-Jul	.	1993-Jan	.
	Core inflation	.	.	1991-Jan	.	.	1983-Mar	.	.
	Energy inflation	.	.	.	.	.	1986-Feb	.	.
	Food inflation	.	.	1991-Feb	.	.	1987-Sep	.	.
	Aggregate-foreign components	.	.	1990-Aug	.	.	1980-Feb	1993-Jan	.
Switzerland	Aggregate inflation	.	1984-Oct	.	.	.	1983-Feb	.	.
	Core inflation	.	.	1993-May	.	1977-Apr		.	.
	Energy inflation	.	1982-Feb	1991-Jun	.	.	1983-Feb	.	.
	Food inflation	.	.	.	.	.	1985-Sep	.	.
	Aggregate-foreign components	.	.	1993-Sep	.	.	1983-Feb	.	.
UK	Aggregate inflation	.	1980-May		.	.	1982-Jul	.	.
	Core inflation	.	.	1990-Sep	.	.	.	1990-May	.
	Energy inflation	1977-Jun	.	.	.	.	.	.	.
	Food inflation	.	1980-Apr	.	.	1977-Mar	.	.	.
	Aggregate-foreign components	1977-May	1984-May	1990-Dec	.	.	1982-Jul	.	.
US	Aggregate inflation	1977-Nov	.	1990-Oct	2003-Feb	.	1983-May	.	
	Core inflation	.	1986-Jun	1993-May	2002-Oct	1977-Jul	1983-Dec	.	.
	Energy inflation	.	1980-Jan 1986-Jul	1993-Jul	.	.	.	.	2000-Jun
	Food inflation	.	.	.	2004-Aug	1978-Aug	.	1995-Jul	.
	Aggregate-foreign components	.	.	1990-Oct	2004-Jul	.	1983-May	.	.

Notes: Estimated dates of structural breaks are shown for the model in equation (6) applied separately to aggregate, core, energy and food inflation. In addition, the row labelled aggregate-foreign components shows estimated dates of structural breaks for the model in equation (9).

Table A.3. Detailed results for coefficient break tests applied to aggregate inflation

Country	WDmax test		Sequential *F-tests*							
	Statistic	p-value	SupF(2|1)	p-value	SupF(3|2)	p-value	SupF(4|3)	p-value	SupF(5|4)	p-value
Austria	21.597	0.002	15.411	0.058	4.952	0.999	9.518	0.705	0.723	1.000
Canada	67.341	0.000	17.583	0.222	9.060	0.999	12.027	0.952	0.000	1.000
Denmark	34.735	0.000	18.457	0.092	5.980	1.000	0.000	1.000	0.000	1.000
Finland	29.233	0.000	16.692	0.166	15.294	0.357	7.876	1.000	0.000	1.000
France	82.217	0.000	23.318	0.015	20.216	0.074	5.206	1.000	0.000	1.000
Germany	50.601	0.000	46.740	0.000	3.643	1.000	3.471	1.000	0.000	1.000
Italy	36.071	0.000	23.448	0.031	22.157	0.075	10.221	0.998	0.000	1.000
Japan	27.923	0.006	11.397	0.923	18.024	0.424	10.435	1.000	10.403	1.000
Netherlands	28.030	0.001	17.027	0.149	7.154	1.000	10.054	0.971	0.000	1.000
Sweden	28.299	0.001	11.491	0.640	7.892	0.998	0.000	1.000	0.000	1.000
Switzerland	38.885	0.000	14.560	0.313	11.634	0.771	11.032	0.914	0.000	1.000
UK	27.609	0.001	16.934	0.154	17.098	0.212	15.094	0.462	0.000	1.000
US	96.590	0.000	17.890	0.021	20.776	0.009	0.000	1.000	0.000	1.000

Notes: Results are shown for structural break test statistics applied to the coefficients of model in equation (6) for aggregate inflation using the iterative procedure of Section 3.2; *p*-values are obtained using the method of Hall and Sakkas (2013).

Table A.4. Detailed results for variance break tests applied to aggregate inflation

Country	WDmax test		Sequential *F-tests*							
	Statistic	p-value	SupF(2|1)	p-value	SupF(3|2)	p-value	SupF(4|3)	p-value	SupF(5|4)	p-value
Austria	21.365	0.000	9.712	0.063	4.051	0.760	2.189	1.000	1.026	1.000
Canada	13.419	0.004	3.163	0.787	4.226	0.728	2.957	0.978	0.760	1.000
Denmark	34.775	0.000	22.886	0.000	2.320	0.993	0.686	1.000	0.000	1.000
Finland	50.690	0.000	7.542	0.162	1.257	1.000	0.000	1.000	0.000	1.000
France	17.090	0.000	1.074	1.000	3.860	0.795	3.232	0.956	0.000	1.000
Germany	17.366	0.000	7.382	0.174	2.571	0.978	0.112	1.000	0.000	1.000
Italy	66.806	0.000	12.101	0.021	3.673	0.829	0.000	1.000	0.000	1.000
Japan	83.061	0.000	48.147	0.000	14.654	0.009	2.050	1.000	0.000	1.000
Netherlands	18.560	0.000	5.404	0.382	2.485	0.984	2.965	0.978	0.000	1.000
Sweden	34.915	0.000	16.260	0.003	2.711	0.965	0.961	1.000	0.000	1.000
Switzerland	37.824	0.000	5.842	0.323	2.886	0.947	0.000	1.000	0.000	1.000
UK	61.482	0.000	7.048	0.200	9.268	0.114	0.893	1.000	0.000	1.000
US	31.606	0.000	10.069	0.053	4.352	0.704	0.018	1.000	0.000	1.000

Notes: As for Table A.3, except that tests are applied to the variance.

Table A.5. Estimates within regimes for domestic/foreign inflation models

Country	Series	Contemporaneous Foreign Inflation				Lagged Foreign Inflation				Mean Inflation				Persistence				Volatility			
		R1	R2	R3	R4	R1	R2	R3	R4	R1	R2	R3	R4	R1	R2	R3	R4	R1	R2	R3	R4
Austria	Aggregate	0.42	0.74			0.30	0.01			0.33	0.18			NA	NA			0.07	0.03		
	Core	0.26	0.11			0.24	0.30			0.63	0.23			0.46	0.23			0.08	0.03		
	Energy	0.14	0.21	0.80		0.03	0.29	0.15		0.63	0.24	0.32		0.24	-0.04	0.07		1.05			
	Food	0.76				NA				0.24				NA				0.35	0.17		
Canada	Aggregate	0.23	0.68			0.29	0.07			0.56	0.17			0.38	0.00			0.09	0.05		
	Core	-0.08	0.16			0.50	-0.05			0.53	0.15			0.52	0.19			0.07	0.03		
	Energy	0.18	0.82	0.41		0.15	0.03	0.02		0.71	0.34	0.19		0.16	0.07	-0.08		1.13	3.17		
	Food	0.62	0.87			0.36	0.02			0.66	0.19			0.11	0.01			0.74	0.47	0.21	
Denmark	Aggregate	0.37	0.47			NA	NA			0.79	0.24			0.40	0.14			0.19	0.08	0.02	
	Core	0.36	0.06			NA	NA			0.66	0.19			0.33	0.53			0.09	0.02		
	Energy	0.56	0.72			0.53	-0.01			1.33	0.28			NA	NA			1.76	0.58		
	Food	0.54	0.36	0.95		0.31	-0.01	0.33		0.76	0.17	0.19		NA	NA	NA		0.43	0.16		
Finland	Aggregate	0.60	0.71			0.17	0.12			0.71	0.17			0.24	0.04			0.20	0.04		
	Core	0.37	0.11			0.43	-0.07			0.69	0.14			0.23	0.36			0.20	0.08	0.03	
	Energy	0.11	0.10	1.13		NA	NA	NA		1.12	0.23	0.34		NA	NA	NA		2.28			
	Food	-0.23	0.71			0.80	0.00			0.96	0.27			0.18	0.15			0.85	0.23		
France	Aggregate	0.32	0.47	0.88		NA	NA	NA		0.73	0.23	0.13		0.65	0.15	0.15		0.04	0.01		
	Core	0.31	0.18			0.09	0.13			0.70	0.15			0.60	0.43			0.04	0.01	0.01	
	Energy	0.23	0.85			NA	NA			0.98	0.19			0.22	0.12			1.11			
	Food	0.42	0.80			NA	NA			0.80	0.17			0.36	0.15			0.06	0.10		
Germany	Aggregate	0.13	1.00	0.72		NA	NA	NA		0.41	0.24	0.16		0.26	-0.15	0.09		0.04	0.02		
	Core	-0.05	0.01			0.09	0.23			0.45	0.19			0.45	0.37			0.04			
	Energy	0.22	1.00	0.68		NA	NA	NA		0.74	0.22	0.40		NA	NA	NA		1.22	0.67		
	Food	0.55	0.69	0.89	1.34	NA	NA	NA	NA	0.27	0.15	0.03	0.21	NA	NA	NA	NA	0.11	0.05	0.12	
Italy	Aggregate	0.07	0.20	0.39		0.57	0.12	0.13		1.06	0.44	0.18		0.64	0.39	0.33		0.16	0.04	0.01	
	Core	0.05	0.40			0.42	-0.11			0.68	0.15			0.73	0.26			0.19	0.05	0.02	
	Energy	0.06	0.23	0.51		0.59	0.16	0.22		0.90	0.79	0.31		0.32	0.08	0.11		2.55			
	Food	0.27	0.25			0.23	0.06			0.78	0.18			0.67	0.51			0.24	0.04		

Table A.5 continued

Country	Series	Contemporaneous Foreign Inflation				Lagged Foreign Inflation				Mean Inflation				Persistence				Volatility			
		R1	R2	R3	R4	R1	R2	R3	R4	R1	R2	R3	R4	R1	R2	R3	R4	R1	R2	R3	R4
Japan	Aggregate	0.31	0.29			0.63	0.21			0.84	0.11			0.34	0.23			0.40	0.13	0.08	0.03
	Core	0.13	0.12			NA	NA			0.37	-0.02			0.75	-0.03			0.26	0.05	0.01	
	Energy	0.04	0.08			0.07	0.14			0.20	0.26			0.43	0.20			0.38	0.06	0.25	1.05
	Food	0.34				0.34				0.21				0.24				1.22	0.47		
Netherlands	Aggregate	0.45	0.39			NA	NA			0.42	0.18			0.57	0.30			0.06	0.02		
	Core	0.24				NA				0.29				0.68				0.04			
	Energy	0.09	0.56			0.07	0.07			0.87	0.22			NA	NA			0.22	1.47	0.80	
	Food	0.62	0.76	0.85		NA	NA	NA		0.58	0.15	0.12		0.12	-0.06	0.13		0.12	0.21	0.08	
Sweden	Aggregate	0.42	0.81			NA	NA			0.67	0.15			0.34	0.14			0.28	0.16	0.04	
	Core	0.06	0.01			0.28	0.66			0.63	0.14			0.40	0.26			0.19	0.06		
	Energy	0.72				NA				0.62				NA				2.48	1.25		
	Food	0.29	0.95			NA	NA			0.73	0.09			0.27	0.05			0.44	0.16		
Switzerland	Aggregate	0.28	0.87			NA	NA			0.40	0.13			0.58	0.29			0.12	0.04		
	Core	0.06	0.04			NA	NA			0.36	0.07			0.83	0.27			0.08	0.02		
	Energy	0.95	2.49	1.33		NA	NA	NA		0.64	-0.17	0.22		NA	NA	NA		3.43	1.61		
	Food	0.69				NA				0.19				0.14				0.43	0.15		
UK	Aggregate	0.56	0.59			NA	NA			1.03	0.31			0.71	0.40			0.12	0.03		
	Core	0.29	0.02			0.20	0.20			0.78	0.18			0.59	0.39			0.11	0.03		
	Energy	0.16	0.55			NA	NA			1.11	0.50			0.38	0.29			1.05			
	Food	0.43	0.55			0.35	0.16			1.16	0.31			0.38	0.16			0.63	0.21		
US	Aggregate	0.63	0.46	0.37	1.36	NA	NA	NA	NA	0.52	0.49	0.22	0.19	0.04	0.49	0.11	0.12	0.07	0.02		
	Core	-0.11	0.37	0.04	0.00	0.29	-0.04	0.04	-0.05	0.54	0.36	0.22	0.16	0.74	0.01	-0.14	0.67	0.03	0.05	0.01	
	Energy	0.29	0.00	1.26		NA	NA	NA		0.74	0.80	0.26		0.64	0.44	0.20		0.86	2.45		
	Food	0.44	0.61			NA	NA			0.37	0.22			0.27	0.22			0.44	0.19	0.07	

Notes: Estimated values within regimes (indicated by the prefix R) for the models in equation (6) estimated separately for aggregate, core, energy and food inflation. Values shown for lagged foreign inflation and persistence are shown as the sums of the relevant estimated coefficients. The mean is calculated as the sample mean of domestic inflation given the estimated break dates, while volatility refers to the estimated disturbance variance in equation (6). NA means not applicable, since no corresponding values are estimated.